# Risk Appetite and the Pricing of Sovereign Risk

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**Abstract**

In modern markets, the bonds of riskier sovereigns are more sensitive to shifts in global risk appetite. The greater sensitivity could reflect the revaluation of risk by the intermediaries that price sovereign debt or news about borrowers' repayment capacity. Using sovereign bonds traded on the London Stock Exchange from 1870 to 1907, this paper tests whether the same pattern holds outside the sample in which it was first found and which account explains it. Borrowers with a history of default load about four times as much as other borrowers on a credit factor built from British corporate bond returns and respond more than twice as much to surprises in Bank Rate, which London's money-market rates followed. In a model in which intermediaries apply a common price of risk to all sovereign claims, a change in that price moves risky claims more but does not forecast export earnings; news about repayment capacity does both. Innovations in the credit factor have no detectable relation to borrowers' export growth, whereas innovations in British equity returns are positively related to export growth in the same year and the next. This evidence is consistent with the greater exposure of riskier sovereigns to shifts in risk appetite being a durable feature of intermediated sovereign lending, and changes in the price of risk account for part of that exposure.

**Keywords:** Global financial cycle; sovereign risk premia; intermediary asset pricing; monetary policy; London Stock Exchange.
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# 1 Introduction

In modern sovereign-debt markets, a contraction in the risk-bearing capacity of global intermediaries widens the spreads of speculative-grade sovereigns three to four times as much as those of investment-grade sovereigns (Gilchrist et al., 2022). Sovereign credit spreads also share a common factor, tied to U.S. equity and volatility risk premia, that accounts for most of their variation (Longstaff et al., 2011). The greater exposure of risky borrowers could reflect a revaluation of risk or news about export earnings and defaults. It could also come from intermediaries that price sovereign debt or from the institutions now surrounding sovereign borrowing, such as dollar funding, investment rules tied to credit ratings, and an official lender to sovereigns (Gilchrist et al., 2019; Chen et al., 2014; Corsetti et al., 2006). Because the modern evidence comes from a period in which these features operate together, whether that greater exposure is a durable feature of intermediated sovereign lending or a feature specific to that period is an open question.

In this paper, I test whether risky borrowers were also more exposed to shifts in risk appetite in the London market between 1870 and 1907, an earlier era of intermediated sovereign lending without rating agencies or an official lender to sovereigns. Because these data played no part in establishing the modern finding, they provide an independent test of whether intermediated lending ranks borrowers' exposure to risk appetite the same way, as other historical data have for patterns in equity returns (Grossman and Shore, 2006; Linnainmaa and Roberts, 2018). Using a hand-collected panel of the sterling bonds of 39 sovereign borrowers, observed at four-week intervals, I find that borrowers with a history of default load about four times as much as other borrowers on a British corporate credit factor and respond more than twice as much to unanticipated changes in Bank Rate, the Bank of England's policy rate, which London money-market rates followed. Innovations in the credit factor have no detectable relation to borrowers' subsequent export growth, whereas innovations in British equity returns, which reflect trade shocks, are positively related to export growth in the same year and the next. The larger loadings of risky borrowers on the credit factor thus point to a revaluation of risk rather than news about export earnings.

A model in which intermediaries price sovereign debt organizes the evidence. In the model, shared money-market funding gives every sterling credit claim a common price of risk. A claim's loading on innovations in that price is its discounted duration times its systematic default risk, so claims on risky borrowers load more. If defaults are independent of the price of risk, as the model assumes, an innovation in that price revalues required premia without forecasting defaults or borrowers' export earnings, and, when the price of risk reverts quickly relative to the timing of a claim's cash flows, its effect on prices largely reverses as expected returns adjust.

News about cash flows would produce the same ranking. Claims on risky borrowers would also move more on news about default losses or about shared fundamentals, such as the prices of the primary commodities most borrowers exported. Such news, however, would be followed by changes in export earnings or defaults, and its effect on prices would persist. British equity returns reflect both the price of risk and trade shocks, which affected British firms and the borrowers alike.

The ranking of loadings therefore cannot separate a change in risk premia from a cash-flow shock, but the difference in what follows an innovation can. Export earnings should follow cash-flow news, but not a change in the price of risk. For borrowers equally exposed to the price of risk, only cash-flow news should make loadings rise with dependence on commodity exports.

Riskier sovereign borrowers load more heavily on the British corporate credit factor and respond more to surprises in Bank Rate. Borrowers with a history of default load 3.7 times as much on the corporate credit factor as borrowers without one, and about four times as much on the equity returns of British banks. That classification uses only information available to investors at the time. A classification of creditworthiness published in 1910 gives a ratio of 4.2, and both classifications rank borrowers as modern estimates do. The ranking holds at the bond level, at smaller ratios among borrowers not in default, and across checks for stale price quotations and the attrition of securities from the panel.

The same ranking appears in the response to a change in intermediaries' cost of funding. Among borrowers not in default, the ratio of responses is 1.6 to 2.0, although it is less precisely estimated. Bonds of the safest borrowers move with the consol, and those of risky borrowers about twice as much. Jordà et al. (2019) find no effect of Bank of England rate changes on risk appetite before 1914 in the equity markets of 17 advanced economies; in the foreign government bonds traded in London, where much of the period's cross-border lending was intermediated, the response is ranked by borrower risk. Claims backed by a guarantee—colonial bonds and the Ottoman loan of 1855—load like those of safe borrowers, as a lower systematic default risk implies. Both results are consistent with the model's funding mechanism, although on their own they do not distinguish a revaluation of risk from news about cash flows.

Adding the British equity market factor to the regression does not distinguish the two accounts either. The class difference on the corporate factor is largely common to British credit and equity prices, so with both factors in the same regression it is no longer distinguishable from zero; in the model, the overlap between the risk-appetite and trade-shock components of the two factors produces this attenuation even when risky borrowers bear more systematic default risk.

Several pieces of evidence point to a change in the price of risk, rather than news about fundamentals, as the source of the larger loadings of risky borrowers on the corporate credit factor. First, the class difference in responses to the equity innovation persists for a year; the difference in responses to the credit innovation is no longer distinguishable from zero after a quarter, although the one-year estimates on the credit innovation are too imprecise to show a reversal. Second, the export test detects trade shocks in the equity innovations but finds none in the credit innovations over the three years beginning with the year of the innovation. Third, within a borrower class, dependence on commodity exports raises the loading on the equity innovation by 0.06 to 0.07 (0.03) per standard deviation of the export share but not the loading on the credit innovation. Fourth, controlling for a factor built from the returns of London-traded commodity producers leaves the class difference unchanged.

The default record gives two further tests of the model. The credit innovation has no detectable

forecasting power for defaults, as the model implies for a change in the price of risk. With defaults arriving at about one a year, however, neither does the market innovation, so the test checks the model's prediction rather than distinguishing the two accounts. According to the point estimates, expected losses account for between a third and a half of the implied carry of risky borrowers and less for the others, although the confidence bands are wide.

The evidence is thus consistent with the greater exposure of risky borrowers being a durable feature of intermediated sovereign lending, and the revaluation of risk by the intermediaries that priced sovereign debt, rather than news about export earnings, accounts for part of it.

## 2 Related research

This paper relates to research on the global financial cycle and on the pricing of sovereign debt before 1914. Modern research attributes much of the common variation in sovereign spreads and risky asset prices to global risk appetite and to the intermediaries that bear risk (Longstaff et al., 2011; Rey, 2013; Bruno and Shin, 2015; Miranda-Agrippino and Rey, 2020). Its estimates of how that exposure differs across borrowers, which find speculative-grade sovereigns the most exposed, come from emerging-market dollar bonds since the mid-1990s (Borri and Verdelhan, 2010; Gilchrist et al., 2022). I test whether intermediated lending produced the same ranking in an earlier era, in data that played no part in establishing it. The main differences from the modern estimates lie in measurement: holding-period returns rather than spreads, and borrower classes based on default history and a contemporary assessment of creditworthiness rather than credit ratings.

Research on the market before 1914 attributes its comovement to news and fundamentals rather than to risk appetite. Mauro et al. (2002, 2006) conclude that the period's sharp spread movements were dominated by country-specific news, such as wars and political violence, and Bordo and Murshid (2006) that its comovement in spread levels may reflect shared long-run trends in fundamentals rather than short-run responses to financial distress. Jordà et al. (2019) find that national equity prices responded to Bank of England rate changes as a risk-neutral investor's would, although cross-border finance was then concentrated in public bonds and railways rather than in equity markets. I examine instead the common variation in four-week sovereign returns after controlling for interest-rate changes, where short-run financial comovement is easier to detect than in spreads, and the response of those returns to Bank Rate surprises by borrower class. The findings are compatible: a first principal component that accounts for a fifth of issuer-level return variation leaves most of it to country-specific news, which can dominate large movements in spreads even when a common price of risk moves all prices together.

Research on the London market shows that investors assessed repayment capacity and the standing of the issuing houses (Flandreau, 2003; Flandreau and Zumer, 2004; Flandreau and Flores, 2009), that a default lowered the prices of other borrowers' bonds issued through the same house (Indarte, forthcoming), and that Empire membership lowered borrowing costs (Ferguson

and Schularick, 2006; Accominotti et al., 2011). The last two bear on tests in this paper: contagion through issuing houses on the test that excludes periods with default entries, and implicit guarantees on the comparison of colonial bonds with safe sovereigns. Meyer et al. (2022) supply the default and haircut record behind the default-history classification and the expected-loss estimates. In the same market, Alquist (2010) shows that liquidity risk is priced in sovereign bonds, and Alquist and Chabot (2011) that gold adherence, which Bordo and Rockoff (1996) argue lowered borrowing costs as a seal of approval, did not lower risk-adjusted returns; controlling for the illiquidity factor or for gold adherence leaves the class difference identified here in place.

# 3 Intermediaries and the common factor

A simple model of an intermediary that prices sovereign debt serves two purposes (see also Gilchrist et al., 2022). It shows why claims on risky borrowers load more on innovations in a common price of risk, and it shows what distinguishes such a revaluation of risk from news about borrowers' repayment capacity, which would produce the same ranking.

## 3.1 The model

A representative London intermediary prices every sterling credit claim. Let $\lambda_t$ denote the intermediary's price of risk; risk appetite is its inverse, so a high price of risk represents low risk appetite. It follows $\lambda_{t+1} = \bar{\lambda} + \rho(\lambda_t - \bar{\lambda}) - u_{t+1}$ with $0 \leq \rho < 1$ and $E_t[u_{t+1}] = 0$, so that a positive innovation $u_{t+1}$ lowers the price of risk and raises risk appetite. The empirical analysis uses the corporate credit factor as a proxy for $u$.

Claim $i$ bears a quantity of risk $q_i$, the conditional covariance of its default loss with the loss on the rest of the intermediary's book. This covariance is the systematic part of the claim's default risk; the book is diversified, so the claim's own weight in it is negligible. Holding the dependence of a claim's default losses on the rest of the book fixed, the covariance is proportional to the claim's loss given default. A default loss independent of the losses on the rest of the book contributes no systematic risk, however large the claim's expected loss.

The intermediary requires claim $i$ to earn a premium of $\lambda_t q_i$ over the bill each period, so the claim's one-period excess return is

$$r^e_{i,t+1} \;=\; \lambda_t \, q_i \;+\; \beta_i \, u_{t+1} \;+\; d_{i,t+1}, \tag{1}$$

where $d_{i,t+1}$ is the claim's default component. Default arrives with per-period intensity $\pi_i$, on default the claim loses fraction $\ell_i$ of its value, and $d$ is mean zero conditional on time-$t$ information. The claim's carry, $s_{i,t}$, is its promised excess return over the bill, the return it pays in a period with no default and no unexpected revaluation. Carry is the premium plus the actuarial component that compensates for expected loss, and it moves with $\lambda_t$ because the price does. The coefficient $\beta_i$ is the revaluation of the claim when the price of risk moves. Because the claim's price is the

present value of its cash flows at the required premium, $\beta_i = \kappa_i q_i$ to first order, where $\kappa_i \geq 0$ is a $\rho$-discounted duration that depends on the persistence of the price of risk and on the claim's cash-flow weights. It is strictly positive for claims with cash flows beyond the next period. Among claims with the same $\kappa_i$, one with twice the quantity of risk is revalued twice as much by the same innovation. If persistence and cash-flow weights make $\kappa_i$ similar across the sample's bonds, differences in loadings primarily reflect differences in $q_i$. Equation (1) thus states that a claim's excess return is the premium the intermediary requires, plus its revaluation when the price of risk moves, plus its default component. Appendix A derives the propositions.

### 3.2 Assumptions and predictions

Two assumptions underlie the results: a common price of risk and default losses that are independent of innovations in the price of risk. Under a common price of risk, every claim is priced with the same $\lambda_t$, and claims differ only in their quantity of risk $q_i$ and their cash-flow weights. Under independent defaults, each claim's default-arrival and loss-given-default processes are independent of the entire sequence of innovations $\{u_t\}$. Hence, an innovation moves the price of risk without conveying news about default losses. Innovations also convey no news about the path of the bill rate.

The assumptions are less stringent than they may appear. A common price of risk does not require a single holder of every claim.[1] Independent defaults allow default losses to be correlated across borrowers and rule out only their correlation with innovations in the price of risk. The model also takes the price of risk as given rather than deriving it from intermediary capital (He et al., 2017), so it does not depend on a particular account of why risk appetite moves.

Shared access to London money-market funding motivates a common component in intermediaries' risk-bearing capacity.[2] The jobbers who made markets in sovereign bonds financed their inventories with callable loans from the London money market secured on the securities held: "It was credit not capital that was important" (Michie, 1999, pp. 105–111). The rates on this credit, including the contango rate, followed Bank Rate (Warren, 1903). Banks, discount houses, and insurance and trust companies also held securities for short periods at rates set against the bill rate, and most purchases on the Stock Exchange were financed on credit (Michie, 1999, pp. 133–135). which assumes that money-market funding costs set the prices at which all holders trade.

The assumption of a common price of risk delivers the cross-sectional prediction. The assumption of independent defaults delivers the time-series prediction and also keeps the loading free of default-loss news. Because a positive $u$ raises risk appetite and the empirical analysis examines

[1] Paris became a serious competitor late in the century, and the houses originated through partners in several cities, with Rothschild the leading issuing house in both (Flandreau et al., 2010, pp. 55–56 and Table 1). The centers were closely connected: the European foreign-exchange network of 1890–1910 was dense and grew denser (Flandreau and Jobst, 2005).

[2] In the primary market, the issuing houses increasingly took sovereign loans firm and, after 1890, spread the risk of an unfilled subscription across underwriters funded from the same money market (Flandreau et al., 2010, pp. 67–68; Chapman, 1984, pp. 49, 88–89).

responses to risk-off innovations, which are negative values of $u$, a claim with a positive loading loses value when risk appetite contracts.

**Proposition 1.** *Under a common price of risk and independent defaults, the loading of claim $i$'s excess return on $u$ is $\beta_i = \kappa_i q_i$, so for claims $i$ and $j$ with nonzero loadings the loadings are in the ratio $\beta_i/\beta_j = \kappa_i q_i/\kappa_j q_j$. For claims with the same $\rho$-discounted duration $\kappa$, the loadings are in the ratio of the quantities of risk, $\beta_i/\beta_j = q_i/q_j$, and claims with the same $\kappa$ and the same $q$ have the same loading.*

Proposition 1 thus predicts the cross-section of loadings: a contraction in risk appetite revalues every claim in proportion to the systematic risk it bears. The ranking of loadings identifies the ranking of $q_i$ when claims have the same discounted duration. More generally, because $\beta_r/\beta_s = (\kappa_r/\kappa_s)(q_r/q_s)$, a risky claim $r$ that loads more than a safe claim $s$ and has a $\kappa$ no larger must have the larger $q$, and the ratio of loadings is then a lower bound on the ratio of $q$. The weights $(1-\rho^{k-1})/(1-\rho)$ in $\kappa_i$ rise with $k$ for $\rho > 0$, so $\kappa_i$ is smaller when more of a bond's present value comes from earlier cash flows, as it does, for given promised cash flows, with a deep discount or a sinking fund that retires the bond on a schedule.

The default record and the contemporary classification rank borrowers by default risk. Applying Proposition 1 to these classes requires the additional assumption, maintained in the empirical analysis, that riskier classes also have larger $q_i$, so that their default losses covary more strongly with losses on the intermediary's portfolio. The assumption cannot be tested directly, but two facts support it. First, countries that had defaulted before were disproportionately those that defaulted in later worldwide crises (Lindert and Morton, 1989). Second, defaults that clustered around crises in the lending centers brought longer spells and larger losses to investors than isolated ones (Kaminsky and Vega-García, 2016).

**Proposition 2.** *Under independent defaults, (i) $u$ does not forecast default arrivals; (ii) the projection of any claim's unexpected excess return on $u$ is a revaluation of its premium path, $\kappa_i q_i u$, with no revision of expected default losses, and the carry decomposes into the actuarial component and the premium, $s_{i,t} = \pi_i \ell_i + \lambda_t q_i$.*

Defaults in this period reflected borrowers' financial and political conditions, including wars (Mauro et al., 2002), but a squeeze that cut a borrower off from refinancing would make defaults follow risk-off innovations: most Latin American defaults of 1820–1931, including those after the crises of 1873 and 1890, clustered around panics in the financial centers (Kaminsky and Vega-García, 2016). The count regressions test whether defaults followed risk-off innovations in this sample.

Independent defaults also rule out a common shock to borrowers' ability to pay. Most borrowers exported primary commodities (Clemens and Williamson, 2004), whose prices were set in world markets, and a fall in the price of a country's principal export raised the market's expectation that it would devalue (Mitchener and Pina, 2020). A common fall in commodity prices

would thus strain the finances of commodity exporters and raise expected losses most for borrowers nearest default, producing the ranking of Proposition 1 through revisions to expected losses rather than to the premium. The two interpretations differ in what follows the innovation. News about fundamentals is followed by changes in export earnings and, where it anticipated them, by defaults, and among borrowers equally near default it moves the loadings of the more commodity-dependent borrowers more. Under the model, a shift in the price of risk does neither; it moves loadings in proportion to $\kappa_i q_i$.

The empirical analysis uses two returns from the lender's market as indicators of the price of risk, the British corporate factor and the British market factor. Write the corporate factor as $f^c_t = u_t + \nu_t$, where $\nu_t$ is idiosyncratic news about the firms in the corporate portfolio, and the market factor as $f^m_t = \xi u_t + \zeta_t$ with $\xi > 0$, where $\zeta_t$ is a trade shock, news about world trade and commodity demand on which both British firms and the borrowers depended. The innovations $u$, $\nu$, and $\zeta$ are mutually independent, with $\sigma^2_u > 0$ and $\sigma^2_\zeta > 0$, and independent of the default surprise $d$. The news $\zeta$ revises borrowers' expected export earnings and, through them, their repayment capacity, whereas $u$, as before, carries no news about either, so $\zeta$ forecasts borrowers' export earnings and $u$ does not. Claim $i$'s excess return becomes

$$r^e_{i,t+1} \;=\; \lambda_t\, q_i \;+\; \kappa_i q_i\, u_{t+1} \;+\; \chi_i\, \zeta_{t+1} \;+\; d_{i,t+1}, \tag{2}$$

where $\chi_i \geq 0$ is the claim's sensitivity to the news. It is larger for a borrower nearer default, for which the same change in export earnings moves the prospect of repayment more, and for a borrower more dependent on commodity exports. Equation (2) is a local return specification: $\zeta$ may revise expected payments and default losses, but I hold $q_i$ fixed, so the news leaves future required excess returns unchanged, and $d_{i,t+1}$ is the remaining mean-zero default-loss innovation, orthogonal to the contemporaneous factor innovations. The revaluation caused by $\zeta$ is therefore permanent, whereas the revaluation caused by $u$ is offset by lower expected returns as the price of risk reverts.

**Proposition 3.** *Under equation (2), (i) a claim's loading on $f^m$ rises with its cash-flow sensitivity $\chi_i$, whereas its loading on $f^c$ does not depend on $\chi_i$; (ii) while the claim remains outstanding and $q_i$ is fixed, its cumulative response from the impact of an innovation $u_{t+1}$ through the $H$ periods that follow is $q_i u_{t+1}\big[\kappa_i - (1-\rho^H)/(1-\rho)\big]$, in which the adjustment of expected returns offsets the initial revaluation, largely so when the price of risk reverts quickly relative to the timing of the claim's cash flows, whereas its cumulative response to $\zeta_{t+1}$ is $\chi_i \zeta_{t+1}$ at every horizon; (iii) the claim's loading on $f^c$ alone is proportional to $\kappa_i q_i$, and in a regression on $f^c$ and $f^m$ together its coefficient on $f^c$ is proportional to $\kappa_i q_i - \xi \chi_i$.*

Proposition 3 separates the two indicators. The corporate factor ranks claims by $\kappa_i q_i$ alone, so the ratio of loadings on it is the ratio of Proposition 1, which is why the analysis uses it to measure exposure to the price of risk. The market factor ranks claims by both $\kappa_i q_i$ and $\chi_i$. If risky borrowers bear more systematic risk and are also more sensitive to trade shocks, both factors rank them the

same way, but only the market factor's innovations should forecast export earnings, move the returns of commodity-dependent borrowers more, and produce persistent responses; responses to the corporate factor's innovations should largely reverse when the price of risk reverts quickly relative to the timing of the claims' cash flows. With both factors in the regression (part (iii)), the class difference in the coefficient on the corporate factor is proportional to $(\kappa_r q_r - \kappa_s q_s) - \xi(\chi_r - \chi_s)$ and is shrunk further by the variance of the factor's own news $\nu$, so a small conditional difference need not imply a small difference in systematic default risk.

The model thus gives the empirical analysis five predictions. First, risky borrowers load more on the corporate factor than safe borrowers (Proposition 1). Second, under moderate persistence, responses to the corporate factor's innovations largely reverse, whereas responses to the market factor's persist (Proposition 3). Third, the corporate factor's innovations do not forecast borrowers' export earnings, whereas the market factor's do. Fourth, dependence on commodity exports raises loadings on the market factor but not on the corporate factor. Fifth, the corporate factor's innovations do not forecast default entries, and, with a positive premium, expected losses account for only part of carry (Proposition 2). The first prediction holds under either account; the second, third, and fourth distinguish the two accounts.

# 4 Measuring risk appetite and borrower risk

## 4.1 Bond prices and excess returns

The sample is a hand-collected panel of 199 sovereign bonds traded on the London Stock Exchange, issued by 39 sovereign borrowers (Table 1). Because the model's common price of risk is that of London intermediaries, a bond is included if it was serviced in London, in sterling or at a stated rate of exchange, and its issuer's debt carried no metropolitan guarantee; the one loan with a foreign guarantee, the Ottoman loan of 1855, is kept and treated as a backed claim. The 39 borrowers are a cross-section comparable to those of modern asset-pricing studies of sovereign credit, which number 26 in Longstaff et al. (2011), 38 in Augustin and Tédongap (2016), and 43 in Andrade et al. (2023).

Prices are bid and ask quotations collected every four weeks (28 days) from the Exchange's official quotation lists as printed in the *Money Market Review* and supplemented from *The Economist*. The collection runs from 1866 through 1907, the last year of the *Money Market Review*'s detailed quotations. Because coverage before 1870 is sparse, the analysis window runs from January 1870 through December 1907, covering 495 four-week periods. Appendix B describes the sources and construction of the panel.

A bond's net return over a four-week period, $r_{i,t+1}$, satisfies

$$1 + r_{i,t+1} = \frac{P_{i,t+1} + C_{i,t+1}}{P_{i,t}}, \tag{3}$$

where $P_{i,t}$ is the midpoint of the bid and ask when both are quoted at both ends of the period,

and otherwise the same quoted side at both ends, and $C_{i,t+1}$ is the coupon paid in the period. Bond-level returns are aggregated to issuer-level portfolios with lagged-market-value weights, with 12,666 issuer-periods in all.

The London banker's bill rate, expressed as a four-week net return $r_{f,t}$, is the risk-free rate. Every return in the paper is an excess return over the bill rate known at the start of the period, $r^e_{i,t+1} = r_{i,t+1} - r_{f,t}$. The bill rate moved with Bank Rate, and Bank Rate changed 273 times between 1870 and 1907, so a money-market tightening enters the excess returns of every bond. For that reason, I control for changes in the bill rate and the consol's current yield, measured as its coupon divided by its mid quotation.

Holding-period returns permit comparisons across bonds whose redemption options and sinking funds make yields to maturity depend on assumptions about future prices (Mauro et al., 2002; Alquist and Chabot, 2011). A redemption option let the borrower repay the bond at par between stated dates, and a sinking fund obliged it to retire a fixed share of the debt each year, by buying bonds in the market or drawing them at par, so the timing of principal depended on the path of prices. Holding-period returns need no assumption about redemption and include the capital gain or loss on a bond that trades away from par. Over a long sample, their average also estimates expected returns, because a realized return differs from the expected return by a forecast error with mean zero. With 39 defaults by 21 of the 39 borrowers between 1870 and 1907, the sample contains the defaults investors feared, which modern samples, with few sovereign defaults, do not. Because the returns are those of bonds that remain outstanding, they omit the gain on a bond drawn at par while trading below it, which understates mean returns on deeply discounted bonds.

Missing and stale quotations may affect the estimated difference between safe and risky borrowers in opposite directions. First, price quotations can lapse during distress, so a return is more likely to be missing in the periods in which a risky borrower's price collapsed, which would flatten the cross-section of loadings. Second, the safest bonds tend to be quoted the least often. A stale quotation delays a safe bond's response, and the safe bond's price catches up later, so the risky-minus-safe difference is steeper on impact and flatter once the responses are cumulated. The difference survives equal weighting and two treatments of the quotation gaps. To address stale quotations among safe bonds, the empirical analysis reports two further measures alongside the impact loadings: the return cumulated over the periods after the innovation, and the loadings on the innovation at $t$, $t-1$, and $t-2$ summed.

### 4.2 The corporate and market factors

The corporate factor is the excess return over the bill on the value-weighted portfolio of British corporate bonds quoted on the London Stock Exchange (Chabot and Kurz, 2010), the counterpart of the default factor in bond factor models (Fama and French, 1993; Elton et al., 1995), measured over the bill rather than over government bonds. Between 1870 and 1907, the portfolio consisted mainly of railway debentures, which made up most of the traded domestic corporate debt of the

period (Coyle and Turner, 2013), so news specific to railways could move the factor.

The rationale for the corporate factor is that British corporate bonds were sterling credit claims traded in the same market and financed from the same money market as sovereign bonds: banks lent to Stock Exchange brokers and jobbers against securities, among them British railway and foreign government issues (Michie, 1999, pp. 133–135). Under a common price of risk, their excess returns therefore move with $u$. The factor's role in the analysis is analogous to that of the excess bond premium, although it is a portfolio return rather than a spread residual. Here the rate changes control for the term component, whereas the excess bond premium removes it by construction.

The credit innovations are the residuals $\hat{u}_t$ from

$$f_t \;=\; \phi_0 + \phi_1 f_{t-1} + \phi_2 f_{t-2} + \boldsymbol{\phi}'\mathbf{z}_t + u_t, \tag{4}$$

where $f_t$ is the corporate factor and $\mathbf{z}_t = (\Delta r_t^{BR}, \Delta r_{t-1}^{BR}, \Delta r_t^{b}, \Delta y_t^{consol}, f_{t-1}^{m}, \sigma_{t-1}^{m})'$ collects the controls: the current and lagged four-week changes in Bank Rate, $r_t^{BR}$; the changes in the bill rate and the consol yield; and the lagged British equity-market excess return and its realized volatility. The residual $\hat{u}_t$, the credit innovation, is the empirical counterpart of the innovation $u_t$ in equation (1); it is measured in percentage points of four-week excess return, and its standard deviation is 1.02. The controls explain 16% of the factor's variance. The residual is orthogonal to the included rates and lagged market variables, but may still contain corporate cash-flow and expected-loss news. I thus interpret it as a proxy for changes in risk appetite rather than as a structurally identified shock to intermediary risk-bearing capacity, in the same way that Gilchrist and Zakrajšek (2012) interpret the excess bond premium.

News specific to the corporate portfolio, such as news about railways, enters the factor as the term $\nu$ of the model. To the extent that such news is independent of the borrowers' returns, it attenuates every claim's loading on the factor in the same proportion and leaves the ratio of loadings unchanged. News that the corporate factor shares with the borrowers' fundamentals would not, and the tests of export earnings and commodity prices address it. The ranking is not specific to the railway composition of the portfolio: it also appears in the responses to Bank Rate surprises and in the loadings on a factor built from the equity of the British banks that financed the market.

The British market factor is the value-weighted excess return on the 969 London-listed equities of firms operating in Britain,[3] so that, like the corporate factor, it measures returns in the lender's own market. It is reported alongside the corporate factor throughout. Its innovation is the residual from equation (4) with the market factor in place of the corporate factor. The controls absorb 20% of its variance, and the correlation between the two innovations is 0.35. Realized volatility is not a

[3]Bank Rate changes come from the daily series assembled by the Bank of England. The equity inputs come from 1,808 London-listed firms from 52 countries, with prices from the Friday official list of the London Stock Exchange as printed in the *Money Market Review* and *The Economist*, and dividends and shares outstanding from the *Investors Monthly Manual* and *The Economist* (Chabot and Kurz, 2010).

factor and enters only as a control in the innovation regression.[4]

### 4.3 Classifying borrowers

I classify borrowers in two ways, one from the default record and one from a contemporary assessment of creditworthiness. The default chronology of Meyer et al. (2022) dates each borrower's defaults, and I use it to form a classification that relies only on information available at the time. Its entry dates and haircuts, which the expected-loss estimates also use, agree with the annual reports of the Corporation of Foreign Bondholders, where both record an episode. Because the chronology dates defaults to the month, a borrower is classified as risky from the first four-week period that begins after the month of its first recorded default, and as safe before it; each period's classification uses only the record known at its start, and the class can change during the sample. Of the 39 sovereign issuers, 21 have a history of default by the end of the sample. Because a borrower is classed as safe until its first recorded default, any misclassification places risky borrowers in the safe class and works against finding a difference between the classes.

The second classification follows Powell (1910), who sorted every national and provincial government borrowing on the Stock Exchange into four grades by the stability of its constitutional government, from countries with a settled constitutional order that made political upheaval remote to societies with no organized government at all.[5] The first three grades are the safe, intermediate, and risky classes of the contemporary classification, and the fourth has no issuer in the panel. The contemporary classification covers 35 of the 39 sovereign issuers and, unlike the default record, has an intermediate class. Powell's rule classifies a loan guaranteed by a European government as safe, regardless of the borrower's class; the 1855 Ottoman loan, guaranteed by Britain and France, is an example.

Powell's classification dates from 1910, after the sample ends, and thus incorporates information that investors did not have at any date in the sample. Egypt illustrates the problem. Its classification as safe reflects foreign financial administration from 1876 and British occupation from 1882, circumstances that an investor in 1875, the year before Egypt defaulted, could not have known. For that reason, the analysis emphasizes the classification based on the default record. The contemporary classification shows whether investors of the period, judging creditworthiness by government stability, ranked borrowers the same way.

Table 2 summarizes the factors and the cross-section. The corporate factor's annualized mean is 1.14% (standard deviation 3.99), the market factor's 2.22 (4.87), and the two correlate at 0.45. By default history to date, borrowers without a history of default return 2.0% a year over the bill and those with one 5.3%, with returns three times as volatile. Under the contemporary classification, the 12 safe sovereigns return 2.5% and the 16 risky ones 5.1%, with returns four times as volatile.

[4] Innovations built with the cross-sectional dispersion of returns in place of realized volatility, or with no volatility control, correlate above 0.99 with the baseline innovations.

[5] Powell (1910, pp. 147–148) writes that the first order comprises "those nations (or those component parts of nations) where ... the existence of a strong central and constitutional authority place[s] any prolonged and serious political and social upheaval outside the range of any but the most remote possibility".

# 5 Risk appetite and the cross-section of sovereign returns

## 5.1 The common factor

Sovereign returns share a common component associated with British corporate bond and British equity returns (Table 3). The specification, which the class comparisons also use, is

$$r^{e}_{i,t} \;=\; \alpha_i + \beta\, f_t + \beta_R\,(f_t \times \mathcal{R}_{i,t}) + \theta\,\mathcal{R}_{i,t} + \boldsymbol{\mu}'\mathbf{w}_t + \boldsymbol{\mu}'_R(\mathbf{w}_t \times \mathcal{R}_{i,t}) + \varepsilon_{i,t}, \tag{5}$$

where $r^{e}_{i,t}$ is issuer $i$'s excess return; $f_t$ is the corporate factor; $\mathcal{R}_{i,t}$ equals one if issuer $i$ is in the risky class of the risk classification in use at the start of period $t$ and zero otherwise; $\mathbf{w}_t = (\Delta r^{b}_{t}, \Delta y^{consol}_{t})'$ collects the rate controls, the changes in the bill rate and in the consol current yield; and $\alpha_i$ is an issuer fixed effect. The coefficient $\theta$ is the difference in returns between classes, which the fixed effect absorbs when the classification does not vary over time. The pooled loading is $\beta$ with the $\mathcal{R}_{i,t}$ terms omitted. Here, $\beta$ is the safe-class loading, $\beta + \beta_R$ is the risky-class loading, and $\beta_R$ is the risky-minus-safe difference reported in Tables 4 and 5. At the bond level, $i$ indexes bonds and $\alpha_i$ is a bond fixed effect. Standard errors are clustered by four-week period throughout. Where stated, the credit innovation $\hat{u}_t$ of equation (4) or the British market factor replaces $f_t$.

In the pooled panel with issuer fixed effects and rate controls, the estimated loading on the corporate factor is 0.40. The first principal component of the pairwise correlation matrix accounts for 20% of issuer-level return variation (Table 3). Rotating each issuer's series in time by an independent random number of periods, which preserves each series' properties but removes alignment across issuers, never gives a share above 6.4% in 5,000 draws, so the common component reflects the alignment of returns across issuers in calendar time rather than chance. Removing each issuer's exposure to the rate changes leaves a share of 18.0%, so the common component is not mainly a response to interest rates.

Figure 1 reports sovereign returns around the British banking crises of the published chronologies and after large changes in Bank Rate. Net of the sample drift, the cross-issuer mean excess return cumulates to $-18.0\%$ over the year after the Baring crisis of November 1890 and to $-10.4\%$ after the 1873 crisis. After Bank Rate hikes of at least one percentage point, it cumulates to $-2.7\%$, small by comparison.

## 5.2 Safe and risky borrowers

Risky borrowers have larger estimated loadings on the corporate factor under both classifications (Table 4, columns 1–3; column 3 is estimated on the 100 bonds of the safe and risky classes). Risky borrowers load 3.7 to 4.2 times as much as safe borrowers, differences of $+0.43$ to $+0.48$, and the 95% intervals of the ratios exclude one.[6]

[6] Risky borrowers' returns are also three to four times as volatile as safe borrowers', but the factor and the rate controls explain a smaller share of their within-issuer variance, 2% against 4 to 5% (Appendix Table C.1). Risky returns carry more idiosyncratic as well as more systematic risk, and the loadings measure the systematic part.

The ranking also holds on the British market factor, on which risky borrowers load 3.7 to 5.0 times as much as safe borrowers (Table 4, columns 4–6). With both factors in the regression, the class difference for default history to date is $+0.81$ on the market factor and $+0.09$ (0.13) on the corporate factor (Table 5, column 5), so the difference is largely common to British credit and equity prices. Because overlapping risk-appetite and trade-shock components can attenuate the conditional difference even when risky borrowers bear more systematic default risk (Proposition 3(iii)), the horse race between the corporate and market factors cannot distinguish the two accounts; the export tests are designed to.

A factor built from the equity of the British banks that financed the market gives the same ranking. On the value-weighted excess return of the 51 British banks in the equity panel, risky borrowers load about four times as much as safe borrowers by default history to date, a difference of $+0.42$ (Table 5, column 7). The difference remains, at $+0.27$, with the market factor, rebuilt to exclude the banks, in the same regression (column 8).

Modern estimates rank borrowers the same way: in the single-factor columns of Gilchrist et al.'s (2022) Table 2 and Figure 3, the same construction gives a ratio of 4 on impact and 3 at 12 months. The comparison concerns the ranking, not the magnitude, because the modern estimates are responses of credit spreads, whose relation to returns depends on duration.[7]

Because the loading is a discounted duration times systematic default risk, a larger loading could reflect a longer duration rather than greater risk. A deeper discount and a sinking fund both shorten duration, however, so differences in duration, if anything, work against the ranking found here.

### 5.3 Bank Rate pass-through by class

Surprise changes in Bank Rate repriced sovereign debt in the order of borrower risk: safe borrowers moved with the consol, and borrowers with a history of default about twice as much. A Bank Rate surprise can affect sovereign returns through consol prices and the cost of financing jobbers' inventories.

A change that the money market had already priced should not move prices when it occurs, so I estimate the responses to the surprise component of each change. I predict the change in Bank Rate from the beginning-of-period gap between the banker's bill rate and Bank Rate, the lagged change, and the lagged level, with the bill rate in the role the futures rate plays in Kuttner (2001). The forecast regression is

$$\Delta r_t^{BR} \;=\; \tau_0 + \tau_1 \left(r_{t-1}^{b} - r_{t-1}^{BR}\right) + \tau_2 \, \Delta r_{t-1}^{BR} + \tau_3 \, r_{t-1}^{BR} + e_t, \tag{6}$$

where $r_{t-1}^{b}$ is the bill rate known at the start of period $t$, $r_{t-1}^{BR}$ is Bank Rate on that date, both in percent per year, and $\Delta r_t^{BR}$ is the net change in the period. The residual $\hat{e}_t$ is the surprise. It

[7] With their global-cycle factor in the same regression, the excess-bond-premium coefficients fall to 0.07 and 0.11, neither distinguishable from zero, so the modern pattern too is a property of the shared risk-appetite component.

measures the change in Bank Rate relative to this forecast's information set and is not a structurally identified monetary-policy shock. The forecast accounts for 20% of the variance of the change (Appendix Table C.2).

The pass-through specification is

$$r^{e}_{i,t} \;=\; \alpha_i + \gamma_{c(i)}\,\hat{e}_t + \varepsilon_{i,t}, \tag{7}$$

where $r^{e}_{i,t}$ is bond $i$'s excess return over the four-week period, and $c(i)$ is its class; where the class varies over time, as default history to date does, its indicators enter in level as well. The coefficient $\gamma_{c(i)}$ is the class's return response, in percent per 100 basis points of surprise. The consol's own excess-return response to the surprise, $-0.52\%$, is the benchmark for a sovereign that moves with the consol (Table 6). Because the surprise is estimated, the standard errors are from a moving-block bootstrap over periods that re-estimates the forecast in every draw.

A 100-basis-point surprise lowers the excess returns of bonds whose issuer had a history of default by an estimated 1.22% and those of the other bonds by 0.60%, a difference of $-0.62$ (0.24); under the contemporary classification, the difference is $-0.83$ (0.25) (Table 6, columns 1 and 2). The safe class's response, $-0.57$, is close to the consol's, as investment-grade dollar sovereigns move with Treasuries in the modern data (Gilchrist et al., 2019). The risky class responds the most. A tightening that raises intermediaries' price of risk revalues each claim in proportion to its systematic risk, so the claims with the most systematic risk should fall the most. They do so even though their deeper discounts shorten duration, which works against the ranking found here.

The ranking survives allowing for stale quotations (Dimson sums, columns 5 and 6) and excluding the crisis periods from both the forecast and the response regressions, so that distress in the bill market does not enter the surprise (Appendix Table C.2). With the risky class split, performing claims respond more than safe ones by $-0.58$ (0.21) under the contemporary classification and $-0.34$ (0.20) by default history, and claims in default respond the most (columns 3 and 4).

### 5.4 Guarantees and backed claims

Two comparisons involve backed claims: colonial bonds, which are otherwise outside the sample, and Ottoman loans with government guarantees or dedicated revenue backing (Appendix Table C.3). In the model, support that reduces a claim's systematic default risk, $q_i$, reduces its loading, holding its discounted duration fixed. Colonial bonds, which carried an implicit metropolitan guarantee (Ferguson and Schularick, 2006; Accominotti et al., 2011) and gained trustee eligibility in 1900, load on the corporate factor like the safe sovereigns, a difference of $+0.02$ (0.04), and far less than the risky class.

Comparing bonds issued by the same borrower holds national characteristics fixed while guarantees differ. The Ottoman loan of 1855, which Britain and France guaranteed during the Crimean War (Nash, 1883), loads 0.12 on the corporate factor, close to the safe sovereigns. Its loading is lower by 0.75 (0.32) than that of the 15 Ottoman issues that carried neither a guarantee nor a

pledge of the Egyptian tribute.

### 5.5 Robustness

Apart from the horse race between the corporate and market factors and the split of the risky class below, the class difference in loadings on the corporate factor ranges from $+0.23$ to $+0.52$ across the checks in Table 5 and Appendix Tables C.3 and C.4. It is distinguishable from zero at the 10% level or better in all but one: at the bond level, assigning the realized return across a quotation gap to its first missing period gives $+0.25$ (0.16). The checks cover alternative factor specifications, two-way clustering by issuer and period, dropping one issuer at a time, equal weighting, stale quotations (summing the loadings on current and lagged innovations, as in Dimson, 1979), the exclusion of the year after each British banking crisis and of the periods in which Bank Rate stood at or above 5%, and interactions of the factor with constraints on the executive (Coppedge et al., 2026), log GDP per capita, and adherence to the gold standard. The Baring crisis of November 1890 brought the larger of the two crisis falls in sovereign returns in Figure 1 and was followed by a change in how the issuing houses underwrote sovereign loans, so the ranking could have held on only one side of it. The ratio of risky to safe loadings exceeds one both before and after the crisis.

Claims on borrowers in default trade as claims on a settlement. With the risky class split into periods of debt service and periods within a default episode (Meyer et al., 2022), performing risky borrowers load more than safe borrowers by $+0.24$ (0.13) by default history to date and $+0.29$ (0.15) under the contemporary classification, and borrowers in default load the most (Appendix Table C.3). The ranking thus holds for performing claims, though less precisely.

Liquidity risk, which Alquist (2010) finds priced in these bonds, does not account for the difference. With the sovereign illiquidity factor in the regression, the difference on the credit innovation is essentially unchanged under either risk classification (Table 5, column 6; Appendix Table C.4).

## 6 Price of risk and repayment capacity

The larger loadings of risky borrowers could reflect a premium for bearing the price of risk or news about default losses and repayment capacity. The tests in this section distinguish a revaluation of risk from news about repayment capacity by what follows an innovation.

### 6.1 Return dynamics

The class difference in responses to the market innovation persists for a year, whereas the difference in responses to the credit innovation is no longer distinguishable from zero after the first quarter. Figure 2 reports local projections of the risky-minus-safe difference in cumulative excess returns on the two innovations. On the credit innovation, the difference is $+0.31$ (0.12) on impact, and the one-year estimates are too imprecise to distinguish a fading difference from a persistent

one (Appendix Table C.5). On the market innovation, it is $+0.68$ on impact and $+1.02$ (0.27) at one year: rather than reversing, it is larger at one year than on impact. The model predicts that repricing caused by a change in the price of risk largely reverses as expected returns adjust when the price of risk reverts quickly relative to the timing of a claim's cash flows, whereas repricing caused by trade shocks persists (Proposition 3(ii)). The persistent difference on the market innovation is what a permanent revision to expected cash flows produces, and it shows that the local projections can detect persistence in these data when it is present. The credit innovation produces no such difference.

### 6.2 Export earnings

Innovations in the credit factor have no detectable relation to borrowers' export growth, whereas innovations in British equity returns do (Table 7). Export earnings provide an annual measure of repayment capacity for the 21 borrowers covered by the export data. To test whether the credit innovation carries news about export earnings, I regress the growth of each borrower's exports over three years on the innovation summed over a calendar year,

$$g_{i,s:s+2} \;=\; a_i + \eta\,\bar{u}_s + \eta_R\,(\bar{u}_s \times \mathcal{R}_{i,s}) + \theta\,\mathcal{R}_{i,s} + e_{i,s}, \tag{8}$$

where $g_{i,s:s+2}$ is the change in the logarithm of borrower $i$'s exports in sterling from the end of year $s-1$ to the end of year $s+2$, $\bar{u}_s$ is the sum of the credit innovations $\hat{u}_t$ over the four-week periods of year $s$, $\mathcal{R}_{i,s}$ is the risky indicator under the risk classification in use at the start of year $s$, $\theta$ is the difference in export growth between classes, which the fixed effect absorbs when the classification does not vary over time, and $a_i$ is a borrower fixed effect. The coefficient $\eta$ is the response of export growth outside the risky class, and $\eta_R$ is the risky-minus-safe difference. Exports are from Obstfeld and Taylor (2003), converted to sterling at their exchange rates. The standard errors are those of Driscoll and Kraay (1998), with a Bartlett window of three years for the overlapping horizon. Under the cash-flow interpretation of the model, a risk-on innovation anticipates higher export earnings, so $\eta > 0$, and $\eta_R > 0$ if the risky class is the more exposed.

For borrowers without a history of default, a one-standard-deviation risk-on credit innovation over a year corresponds to export growth over the three years beginning with the year of the innovation that is lower by an estimated 1.1 percentage points (1.5); at the upper end of the 95% interval, it is higher by 1.8 percentage points. None of the single-year responses (Appendix Table C.5), the difference for borrowers with a history of default, or the estimates under the contemporary classification is distinguishable from zero.

The market innovation shows that the test has the power to detect trade shocks. Under equation (2), its news component revises export earnings, and with the annual sum of the market innovation in place of $\bar{u}_s$ in equation (8), a one-standard-deviation risk-on innovation corresponds to export growth over the three years that is higher by 5.7 percentage points (1.0) under both classifications, spread over the year of the innovation and the two that follow (Appendix Table C.5).

### 6.3 Commodity prices

The cross-section gives a second implication. Under the cash-flow interpretation, a borrower's loading reflects both how much its export earnings depend on commodity prices and how near default it stands. Among borrowers in the same class, and so with similar exposure to the price of risk, the loading should therefore rise with the share of primary commodities in the borrower's exports. To test this idea, I add the interaction of the innovation with that share to equation (5), with the innovation $\hat{u}_t$ in place of the factor $f_t$,

$$\begin{aligned} r^e_{i,t} \;=\; & \alpha_i + \beta\,\hat{u}_t + \beta_R\,(\hat{u}_t \times \mathcal{R}_{i,t}) + \beta_X\,(\hat{u}_t \times X_{i,t}) + \delta\,X_{i,t} + \theta\,\mathcal{R}_{i,t} \\ & + \boldsymbol{\mu}'\mathbf{w}_t + \boldsymbol{\mu}'_R(\mathbf{w}_t \times \mathcal{R}_{i,t}) + \varepsilon_{i,t}, \end{aligned} \tag{9}$$

where $X_{i,t}$ is the primary-product share of borrower $i$'s exports in the previous calendar year (Clemens and Williamson, 2004), standardized across borrower-years so that $\beta_X$ is the change in the loading per standard deviation of the share (20 percentage points). Primary commodities made up more than 80% of exports for 18 of the 23 sovereign borrowers in the panel covered by Clemens and Williamson (2004), averaged between 1870 and 1907. The coefficient $\delta$ absorbs any direct effect of the share on returns, and $\theta$ is the difference in returns between classes when the classification varies over time.

Within a class, dependence on commodity exports does not raise the loading on the credit innovation, as news about exports would. With the class interaction, the estimated coefficient on the export share is $-0.10$ (0.04) for the classification by default history, with the same sign under the contemporary classification (Table 8, columns 1 and 2), whereas Proposition 3(i) implies no relation unless $\kappa_i q_i$ varies with the share within a class. On the market innovation, which reflects trade shocks, the coefficient is positive, 0.06 to 0.07 (0.03) (columns 3 and 4).

Export shares are annual, and I estimate the loadings on four-week returns. The returns of primary-commodity producers outside the European core (the United Kingdom, France, Germany, the Netherlands and Belgium) whose shares traded in London provide a proxy for commodity news measured on the same dates as the bond prices. I construct a commodity factor from the equity panel of Alquist et al. (2022): the value-weighted return on 123 producers of minerals, nitrates, oil, tea, sugar, meat, timber and other primary products operating outside the core and the United States, excluding gold mines, minus the value-weighted return on core equities outside those industries. By average value weight, mining companies make up about two-thirds of the long leg, and nearly half of the long leg operated in Australia and a quarter in South Africa; the short leg is almost entirely British. For each borrower, the long leg omits companies operating in the borrower's country, removing their direct contribution to that borrower's factor. To test whether commodity news accounts for the class difference, I add the factor and its interaction

with the class indicator to equation (5),

$$\begin{aligned} r^{e}_{i,t} \;=\; & \alpha_i + \beta\, f_t + \beta_R\,(f_t \times \mathcal{R}_{i,t}) + \psi\, c_{i,t} + \psi_R\,(c_{i,t} \times \mathcal{R}_{i,t}) + \theta\, \mathcal{R}_{i,t} \\ & + \boldsymbol{\mu}'\mathbf{w}_t + \boldsymbol{\mu}'_R(\mathbf{w}_t \times \mathcal{R}_{i,t}) + \varepsilon_{i,t}, \end{aligned} \tag{10}$$

where $c_{i,t}$ is the commodity factor that leaves out borrower $i$'s country, $\psi$ is the safe class's loading on the commodity factor, $\psi_R$ is the risky-minus-safe difference in that loading, and $\theta$ is the difference in returns between classes when the classification varies over time. If the class difference in loadings on the corporate factor reflected commodity news, the estimate of $\beta_R$ would fall once $c_{i,t}$ enters.

The commodity factor is nearly uncorrelated with the corporate factor, with a correlation of 0.02 over four-week periods, and controlling for it leaves the class difference essentially unchanged under both classifications (Table 8, columns 5 and 6). Risky bonds also load more on the commodity factor itself (columns 5 and 6). That is the cash-flow channel at work: news about commodity prices moves the repayment prospects of borrowers nearer default the most. Because that news is nearly uncorrelated with the corporate factor, it does not account for the larger loadings of risky bonds on the corporate factor.

### 6.4 Default forecasts and default events

The restructuring record of Meyer et al. (2022) contains 39 default entries across 21 issuers within the sample, with episodes running from the recorded default date to the resolution year. Under Proposition 2(i), the credit innovation does not forecast default arrivals. To test this hypothesis, let $N_{t+1:t+h}$ be the number of default entries over the $h$ periods after $t$, and estimate the Poisson regression

$$E\big[N_{t+1:t+h} \mid \hat{u}_t\big] \;=\; \exp\big(\alpha_h + \vartheta_h\, \hat{u}_t\big), \tag{11}$$

by quasi-maximum likelihood, which is consistent for the conditional mean regardless of the count distribution (Gourieroux et al., 1984), with Bartlett standard errors over $h$ lags to account for overlap.

The credit innovation has no statistically detectable relation to default entries at horizons of one to five years (Appendix Table C.2; see also Giesecke et al., 2011). At the upper limit of the 95% interval, a one-standard-deviation risk-off innovation adds at most 0.16 entries over the following year, against an unconditional 1.04.

Nor does the market innovation forecast default entries, although it forecasts export growth. News about repayment capacity can move bond prices without producing a default, and with about one entry a year the count test may lack the power to detect it. The default forecasts are therefore a check on Proposition 2(i) rather than a means of distinguishing a change in the price of risk from news about defaults.

A default entry could also revise expected losses on other borrowers' bonds, for example, when it reflects on the issuing house that brought out both loans. Excluding the four-week periods

that contain the entries, or those periods and the period after each, leaves the class difference essentially unchanged, as does controlling for the number of entries in each period (Appendix Table C.4). Recorded default entries thus account for little of the comovement of sovereign returns. Whatever contagion a default carried to other bonds through a shared issuing house (Indarte, forthcoming), the common component of sovereign returns does not come from default news.

### 6.5 Expected losses and carry

The default record also allows carry to be split into compensation for expected losses and a premium for bearing risk, as in Proposition 2(ii). Expected losses account for about half of implied carry or less in each class under both classifications (Table 9). A bond's carry, the promised excess return in a period with no default or unexpected revaluation, compensates investors for expected default losses and for bearing systematic risk. I use historical defaults, haircuts, and realized returns to estimate these two components, as research on credit default swap premia does (Longstaff et al., 2011), although that research must infer default rates because modern sovereign defaults are few (Ang and Longstaff, 2013). For each class $c$, the expected-loss rate per four-week period is

$$\mathrm{EL}_c \;=\; \pi_c \, \ell_c, \qquad \pi_c = \frac{N_c}{A_c}, \tag{12}$$

where $\pi_c$ is the default-entry hazard rate, the number of entries $N_c$ over the issuer-periods at risk $A_c$, with an issuer at risk when not inside a default episode and its return observed, and $\ell_c$ is the mean present-value haircut of the restructurings that follow those entries in the restructuring record of Meyer et al. (2022).

Table 9 reports the hazard rates, haircuts, expected-loss rates, and mean excess returns by class. The implied carry is the mean realized excess return plus the estimated expected-loss rate, an estimate of the excess return the class's bonds would have earned had no default loss occurred. In Proposition 2, carry equals, to first order, expected default losses plus the risk premium $\lambda_t q_i$, and the mean realized excess return over the quoted issuer-periods at risk estimates the premium. The haircut measures the loss relative to the contractual terms rather than to the pre-default price, which already discounted the default, so the expected-loss rates probably overstate losses in market-value terms.

For borrowers with a history of default, expected losses are 36% of implied carry, so carry is nearly three times expected losses, above the ratio of about two that Giesecke et al. (2011) estimate for U.S. corporate bonds over 1866–2008. For the other borrowers, the share is 24%. The shares are 7%, 20%, and 51% for the safe, intermediate, and risky classes of the contemporary classification.

The bootstrap bands on the shares are wide because the class mean returns are imprecisely estimated. Under the contemporary classification, the 95% band excludes a share of half or more for the safe class; it does not for the safe class by default history, or for the risky class under either classification. In every class, however, the band on the share excludes the case in which expected losses account for all of carry, and the band on the premium excludes zero. Under either risk

classification, carry includes a premium for bearing risk as well as compensation for expected losses.

## 7 Conclusion

In modern sovereign-debt markets, risky borrowers are more exposed to shifts in global risk appetite, and the greater exposure could reflect a revaluation of risk or news about repayment capacity. I test the ranking on the returns of sovereign bonds traded on the London Stock Exchange from 1870 to 1907, a sample that played no part in establishing the original pattern.

Borrowers with a history of default load about four times as much on the corporate credit factor as other borrowers, and their returns are more than twice as sensitive to unanticipated changes in Bank Rate, which London's money-market rates followed. The ranking of borrowers is the one modern estimates also find.

Because the ranking alone cannot distinguish a revaluation of risk from news about cash flows, I examine what follows an innovation. The credit factor's innovations have no detectable relation to borrowers' export growth, whereas the same test detects news in British equity returns, whose innovations are positively related to export growth in the same year and the next. Borrowers' dependence on commodity exports raises their loadings on the equity innovation but not on the credit innovation. That contrast is consistent with the credit factor reflecting changes in required risk premia, while equity returns also reflect trade shocks. The evidence does not precisely establish how much of the greater exposure the price of risk explains: according to the point estimates, expected losses make up between a third and a half of implied carry for risky borrowers and less for the others, but the confidence bands are wide.

The results qualify the view that the comovement of sovereign debt before 1914 reflected mainly country-specific news and shared fundamentals (Mauro et al., 2002; Bordo and Murshid, 2006): in four-week returns, the bonds traded in London shared a financial component whose influence rose with borrower risk. Bank Rate surprises moved the prices of sovereign debt in the order of borrower risk, consistent with the center's policy rate reaching foreign borrowers through the intermediaries that priced their debt. The ranking of sovereign borrowers by their exposure to shifts in risk appetite holds outside the sample in which it was first found, and changes in the price of risk account for part of risky borrowers' greater exposure, as the intermediary account of the global financial cycle implies for today's sovereign-debt markets.

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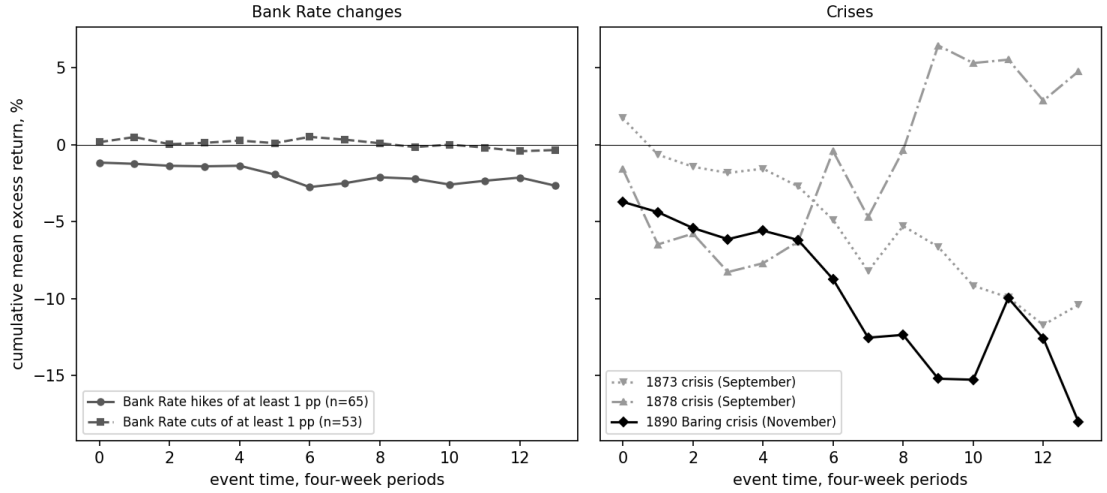
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# Figures and Tables

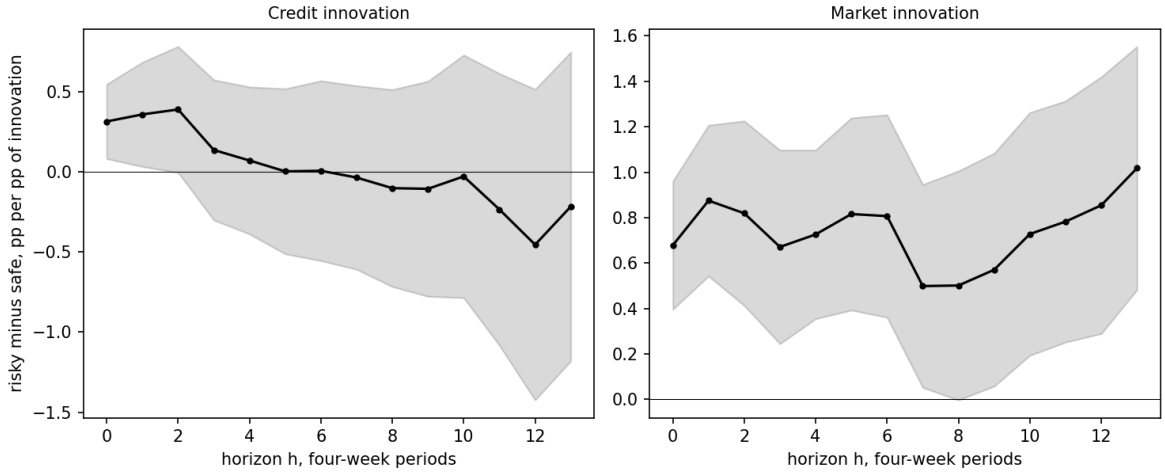
## Figures

Figure 1: Sovereign returns after Bank Rate changes and in crises




*Notes*: Cross-issuer mean excess return over the banker's bill rate, cumulated from the event period and net of the sample drift of 0.30% per period. Left: Bank Rate changes of at least one percentage point in the period. Right: the British banking crises of September 1873, September 1878, and November 1890, each anchored at the first period ending on or after the first of the month. The 1873 month is from Clapham (1945, pp. 293–295), for the year-dated entry of Schularick and Taylor (2012); the 1878 and 1890 months are from Baron et al. (2021).

Figure 2: Risky minus safe: responses to credit and market innovations




*Notes*: Local projections of the risky-minus-safe difference in the mean cumulative excess return from $t$ to $t+h$ on the credit innovation (left) or the market innovation (right) at $t$, without further controls, for $h = 0, \ldots, 13$ four-week periods. Classes follow the contemporary classification (Powell, 1910). The $h = 0$ coefficient is the difference in impact loadings. Bands are 95% Newey–West intervals with $h+1$ lags.

## Tables

Table 1: Borrowers, classifications, and coverage, 1870–1907

| | Class | Bonds | Years | Periods | | Class | Bonds | Years | Periods |
|---|---|---|---|---|---|---|---|---|---|
| Argentina | interm. | 32 | 1870–1907 | 493 | Liberia | risky | 1 | 1902–1907 | 43 |
| Austria-Hungary | safe | 3 | 1873–1907 | 452 | Mexico | interm. | 8 | 1870–1907 | 495 |
| Belgium | safe | 1 | 1874–1898 | 191 | Morocco | uncl. | 1 | 1870–1882 | 82 |
| Brazil | interm. | 14 | 1870–1907 | 488 | Netherlands | safe | 2 | 1870–1907 | 346 |
| Bulgaria | uncl. | 3 | 1890–1907 | 218 | Nicaragua | risky | 1 | 1887–1905 | 234 |
| Chile | interm. | 12 | 1870–1907 | 495 | Norway | safe | 4 | 1882–1907 | 329 |
| China | risky | 10 | 1886–1907 | 286 | Orange Free State | uncl. | 1 | 1885–1895 | 135 |
| Colombia | risky | 4 | 1870–1907 | 422 | Paraguay | risky | 2 | 1873–1907 | 442 |
| Costa Rica | risky | 2 | 1886–1907 | 264 | Peru | interm. | 3 | 1870–1890 | 250 |
| Denmark | safe | 2 | 1895–1907 | 148 | Portugal | risky | 2 | 1870–1907 | 456 |
| Ecuador | risky | 1 | 1870–1900 | 350 | Romania | risky | 3 | 1870–1889 | 250 |
| Egypt | safe | 8 | 1870–1907 | 491 | Russia | interm. | 18 | 1870–1907 | 495 |
| France | safe | 1 | 1872–1875 | 47 | Santo Domingo | risky | 1 | 1870–1890 | 238 |
| Germany | safe | 1 | 1891–1907 | 197 | Spain | risky | 4 | 1870–1907 | 495 |
| Greece | risky | 6 | 1880–1907 | 346 | Sweden | safe | 6 | 1870–1907 | 452 |
| Guatemala | risky | 2 | 1888–1907 | 250 | Turkey | risky | 18 | 1870–1907 | 489 |
| Hawaii | uncl. | 1 | 1888–1898 | 134 | United States | safe | 3 | 1881–1900 | 224 |
| Honduras | risky | 2 | 1870–1907 | 427 | Uruguay | interm. | 6 | 1873–1907 | 442 |
| Italy | safe | 5 | 1870–1907 | 495 | Venezuela | risky | 2 | 1870–1881 | 124 |
| Japan | safe | 3 | 1870–1907 | 451 | | | | | |

*Notes*: The 39 sovereign borrowers in the issuer panel. Class: the contemporary classification (Powell, 1910); interm., intermediate; uncl., not classified. Bonds: bonds with a return in the panel. Years: first and last year quoted. Periods: four-week periods with an issuer return, 12,666 in all. Names are those of the era.

Table 2: Factors, controls, and borrowers, 1870–1907

| | Mean | SD | NW $t$ | Correlations | | | | |
|---|---|---|---|---|---|---|---|---|
| | | | | Corp. | Mkt. | Term | $\Delta$bill | $\Delta y^{consol}$ |
| *Factors and controls* | | | | | | | | |
| British corporate factor | 1.14 | 3.99 | 1.82 | 1.00 | | | | |
| British market factor | 2.22 | 4.87 | 2.60 | 0.45 | 1.00 | | | |
| Consol excess return (term premium) | 0.06 | 4.14 | 0.11 | 0.32 | 0.38 | 1.00 | | |
| Change in bill rate ($\Delta$bill) | 0.00 | 0.06 | 0.30 | −0.13 | −0.07 | −0.00 | 1.00 | |
| Change in consol yield ($\Delta y^{consol}$) | 0.00 | 0.03 | 0.60 | −0.32 | −0.40 | −0.82 | −0.08 | 1.00 |
| British bank factor | 3.43 | 5.84 | 3.93 | 0.25 | 0.44 | 0.29 | −0.00 | −0.19 |
| | Issuers | Bonds | Issuer-periods | Mean | SD | MRT | | |
| *Contemporary classification* | | | | | | | | |
| Safe | 12 | 39 | 3,823 | 2.47 | 7.39 | 2 | | |
| Intermediate | 7 | 93 | 3,158 | 3.84 | 16.64 | 7 | | |
| Risky | 16 | 61 | 5,116 | 5.14 | 29.27 | 15 | | |
| Unclassified | 4 | 6 | 569 | 3.96 | 6.52 | 1 | | |
| *Default history to date* | | | | | | | | |
| No history of default | 23 | 86 | 5,189 | 1.99 | 8.18 | 9 | | |
| History of default | 21 | 130 | 7,477 | 5.33 | 26.23 | 21 | | |
| All issuers | 39 | 199 | 12,666 | 3.96 | 20.82 | 25 | | |

*Notes*: Factors and controls: 495 four-week periods. Means and standard deviations of excess returns over the banker's bill rate are annualized (13 times the four-week mean and $\sqrt{13}$ times the four-week standard deviation, in %), with Newey–West $t$-statistics (five lags). The bill-rate change is the change in the four-week rate and the consol-yield change the change in the annual current yield, both in percentage points per period. Borrowers: annualized mean and standard deviation of issuer-portfolio excess returns, by class. Under default history to date, an issuer-period is in the history-of-default class from the first four-week period beginning after the month of the issuer's first recorded default (Meyer et al., 2022). Issuers and bonds that change class count in both rows. MRT: issuers with an episode in Meyer et al. (2022), including four whose first default came after 1907. Bulgaria, Hawaii, Morocco, and the Orange Free State are unclassified.

Table 3: Comovement in sovereign returns

*Panel A: pooled loading on the corporate factor*

Dependent variable: excess return over the bill rate, $r^e_{i,t}$

| | (1) | (2) |
|---|---|---|
| $\bar{\beta}$ | $0.40^{***}$ | $0.56^{***}$ |
| | (0.09) | (0.10) |
| $\bar{\mu}_b$ | $-0.40$ | |
| | (1.98) | |
| $\bar{\mu}_y$ | $-18.47^{***}$ | |
| | (3.95) | |
| Issuer-periods | 12,648 | 12,666 |
| Within $R^2$ | 0.019 | 0.012 |

*Panel B: share of the first principal component, %*

| | |
|---|---|
| Sovereign excess returns | 20.2 |
| Rotation null: mean / 99th percentile / maximum | 5.3 / 5.8 / 6.4 |
| Residuals on the rate changes | 18.0 |

*Notes*: Panel A: issuer portfolios of the 39 sovereign borrowers, 1870–1907, with issuer fixed effects and one loading on the corporate factor ($\bar{\beta}$) for all classes; column (1) adds the changes in the bill rate ($\bar{\mu}_b$) and the consol yield ($\bar{\mu}_y$). Standard errors clustered by four-week period are in parentheses. *** (**) (*) denotes significance at the 1% (5%) (10%) level. Panel B: share of the largest eigenvalue of the pairwise correlation matrix of the 36 issuers with at least 100 returns. The rotation null rotates each issuer's series in time by an independent random number of periods (5,000 draws).

Table 4: Loadings of sovereign returns on the British factors, by borrower class

Dependent variable: excess return over the bill rate, $r^e_{i,t}$

| | Corporate factor | | | British market factor | | |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) |
| $\beta$ | $0.16^{***}$ | $0.15^{***}$ | $0.14^{***}$ | $0.21^{***}$ | $0.24^{***}$ | $0.28^{***}$ |
| | (0.04) | (0.04) | (0.04) | (0.05) | (0.04) | (0.04) |
| $\beta_R$ | $0.43^{***}$ | $0.48^{***}$ | $0.46^{***}$ | $0.84^{***}$ | $0.81^{***}$ | $0.75^{***}$ |
| | (0.13) | (0.13) | (0.11) | (0.13) | (0.15) | (0.15) |
| $(\beta+\beta_R)/\beta$ | 3.7 | 4.2 | 4.2 | 5.0 | 4.5 | 3.7 |
| 95% interval | [1.8, 5.7] | [2.1, 6.3] | [2.3, 6.1] | [2.9, 7.2] | [3.0, 6.0] | [2.7, 4.7] |
| Classification | History | Contemp. | Contemp. | History | Contemp. | Contemp. |
| Unit | Issuer | Issuer | Bond | Issuer | Issuer | Bond |
| Rate controls | Yes | Yes | Yes | Yes | Yes | Yes |
| Observations | 12,648 | 8,927 | 19,225 | 12,648 | 8,927 | 19,225 |
| Within $R^2$ | 0.023 | 0.019 | 0.024 | 0.045 | 0.036 | 0.043 |

*Notes*: The regression is equation (5), estimated over 1870–1907 (495 four-week periods) on issuer portfolios with issuer fixed effects in columns (1), (2), (4), and (5) and on bonds with bond fixed effects in columns (3) and (6). The indicator $\mathcal{R}$ marks the risky class, by default history to date (History) or the contemporary classification (Contemp.; Powell, 1910), which compares its risky and safe classes. The coefficient $\beta$ is the safe class's loading on the factor in the column heading and $\beta_R$ the risky-minus-safe difference; the 95% interval for $(\beta+\beta_R)/\beta$ is from the delta method. The rate controls are the changes in the bill rate and the consol yield, each interacted with $\mathcal{R}$. Standard errors clustered by four-week period are in parentheses. *** (**) (*) denotes significance at the 1% (5%) (10%) level.

Table 5: The risky-minus-safe difference under alternative factor specifications

Dependent variable: excess return over the bill rate, $r^e_{i,t}$

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
|---|---|---|---|---|---|---|---|---|---|
| $\beta$ | $0.17^{***}$ | $0.13^{***}$ | $0.22^{***}$ | $0.16^{***}$ | $0.09^{**}$ | $0.12^{***}$ | $0.15^{***}$ | $0.11^{***}$ | $0.28^{**}$ |
| | (0.04) | (0.04) | (0.05) | (0.05) | (0.04) | (0.04) | (0.03) | (0.03) | (0.12) |
| $\beta_R$ | $0.43^{***}$ | $0.38^{***}$ | $0.48^{***}$ | $0.43^{***}$ | 0.09 | $0.39^{***}$ | $0.42^{***}$ | $0.27^{**}$ | $0.51^{***}$ |
| | (0.13) | (0.13) | (0.12) | (0.13) | (0.13) | (0.13) | (0.11) | (0.11) | (0.18) |
| $\beta^m$ | | | | | $0.18^{***}$ | | | $0.14^{***}$ | |
| | | | | | (0.05) | | | (0.05) | |
| $\beta^m_R$ | | | | | $0.81^{***}$ | | | $0.64^{***}$ | |
| | | | | | (0.14) | | | (0.12) | |
| $\beta^{\text{il}}$ | | | | | | $0.16^{***}$ | | | |
| | | | | | | (0.04) | | | |
| $\beta^{\text{il}}_R$ | | | | | | 0.03 | | | |
| | | | | | | (0.06) | | | |
| Factor | Corporate | Credit | Credit | Corporate | Corporate | Credit | Bank | Bank | Credit |
| Consol control | Return | Yield | None | Yield | Yield | Yield | Yield | Yield | Yield |
| Bill-rate control | Yes | Yes | No | Yes | Yes | Yes | Yes | Yes | Yes |
| Estimator | Panel | Panel | Panel | Panel | Panel | Panel | Panel | Panel | Dimson |
| Clustering | Period | Period | Period | Two-way | Period | Period | Period | Period | Period |
| Observations | 12,666 | 12,629 | 12,629 | 12,648 | 12,648 | 12,488 | 12,596 | 12,596 | 495 |
| Within $R^2$ | 0.022 | 0.022 | 0.012 | 0.023 | 0.046 | 0.029 | 0.032 | 0.048 | |

*Notes*: The regressions are equation (5) on issuer portfolios with issuer fixed effects, by default history to date, with every factor and control interacted with $\mathcal{R}$. The Factor row names the factor for $\beta$ and $\beta_R$. Column (1) replaces the consol-yield change with the consol excess return. Columns (2), (3), and (6) use the credit innovation of equation (4), built in column (3) without contemporaneous rate controls. Column (5) adds the British market factor ($\beta^m$, $\beta^m_R$) and column (6) the sovereign illiquidity factor ($\beta^{\text{il}}$, $\beta^{\text{il}}_R$; Appendix C). Columns (7) and (8) use the British bank factor (Appendix B); column (8) adds the market factor rebuilt excluding banks ($\beta^m$, $\beta^m_R$). Column (9) reports Dimson sums of the loadings on the credit innovation at $t$, $t-1$, and $t-2$, from regressions of the mean excess return of each class's bonds; the difference and its standard error are from the two class series stacked. Standard errors clustered by four-week period, and in column (4) by issuer and period, are in parentheses. *** (**) (*) denotes significance at the 1% (5%) (10%) level.

Table 6: Bank Rate surprises and sovereign returns, by borrower class

Dependent variable: excess return of a bond over the bill rate, % per 100 basis points of surprise

| | Impact | | | | Dimson sum | |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) |
| $\gamma_S$ | $-0.60^{***}$ | $-0.57^{***}$ | $-0.60^{***}$ | $-0.57^{***}$ | $-0.50^{***}$ | $-0.65^{***}$ |
| | (0.10) | (0.08) | (0.10) | (0.08) | (0.18) | (0.14) |
| $\gamma_I$ | | $-0.83^{***}$ | | $-0.83^{***}$ | | $-1.08^{***}$ |
| | | (0.19) | | (0.19) | | (0.28) |
| $\gamma_R$ | $-1.22^{***}$ | $-1.40^{***}$ | | | $-1.63^{***}$ | $-1.66^{***}$ |
| | (0.24) | (0.28) | | | (0.42) | (0.54) |
| $\gamma_P$ | | | $-0.94^{***}$ | $-1.15^{***}$ | | |
| | | | (0.19) | (0.23) | | |
| $\gamma_D$ | | | $-2.07^{***}$ | $-1.98^{***}$ | | |
| | | | (0.56) | (0.65) | | |
| $\gamma_R - \gamma_S$ | $-0.62^{***}$ | $-0.83^{***}$ | | | | |
| | (0.24) | (0.25) | | | | |
| $\gamma_P - \gamma_S$ | | | $-0.34^{*}$ | $-0.58^{***}$ | | |
| | | | (0.20) | (0.21) | | |
| $\gamma_D - \gamma_S$ | | | $-1.47^{***}$ | $-1.41^{**}$ | | |
| | | | (0.55) | (0.63) | | |
| Classification | History | Contemp. | History | Contemp. | History | Contemp. |
| Bond fixed effects | Yes | Yes | Yes | Yes | Yes | Yes |
| Bond-periods | 36,740 | 35,979 | 36,740 | 35,979 | 36,740 | 35,979 |

*Notes*: The regression is equation (7) on bonds with bond fixed effects, 1870–1907; the surprise is the residual of the forecast regression (6). The coefficients $\gamma_S$, $\gamma_I$, and $\gamma_R$ are the responses of the safe, intermediate, and risky classes; in columns (3) and (4), $\gamma_P$ and $\gamma_D$ are those of risky issuer-periods outside and inside a default episode (Meyer et al., 2022). Columns (1) and (3) classify borrowers by default history to date, which has no intermediate class, and columns (2) and (4) by the contemporary classification (Powell, 1910). Columns (5) and (6) report Dimson sums of the coefficients on the surprise at $t$, $t-1$, and $t-2$. The consol's excess return falls by 0.52 (0.09) per 100 basis points of surprise. Standard errors in parentheses are from a moving-block bootstrap over periods (blocks of 13 periods, 5,000 draws) that re-estimates the forecast regression in each draw. *** (**) (*) denotes significance at the 1% (5%) (10%) level.

Table 7: Credit and market innovations and export growth

Dependent variable: export growth from year $s-1$ to year $s+2$, %

| | Credit innovation | | Market innovation | |
|---|---|---|---|---|
| | (1) | (2) | (3) | (4) |
| $\eta$ | $-1.05$ | $-1.25$ | $5.73^{***}$ | $5.65^{***}$ |
| | (1.45) | (1.51) | (1.04) | (1.08) |
| $\eta_R$ | $-0.48$ | $0.19$ | $-0.25$ | $2.07$ |
| | (0.76) | (1.13) | (0.88) | (1.79) |
| Classification | History | Contemp. | History | Contemp. |
| Issuer-years | 708 | 545 | 708 | 545 |

*Notes*: The regression is equation (8), on annual data for 1870–1909. The innovation is the sum of the four-week innovations over year $s$, scaled to one standard deviation (3.4 and 5.1 percentage points for the credit and market innovations). Exports in sterling are from Obstfeld and Taylor (2003); Brazil's 1887 value, for July to December (Mitchell, 1983, p. 550 and n. 19), is annualized. The sample is the issuer-years with all three years observed, so the three-year response need not equal the sum of the single-year responses in the text. Driscoll–Kraay standard errors are in parentheses. *** (**) (*) denotes significance at the 1% (5%) (10%) level.

Table 8: Commodity exposure and the loadings

Dependent variable: excess return over the bill rate, $r^e_{i,t}$

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| $\beta$ | $0.07$ | $0.10^{**}$ | $0.24^{***}$ | $0.26^{***}$ | $0.16^{***}$ | $0.15^{***}$ |
| | (0.06) | (0.05) | (0.06) | (0.05) | (0.04) | (0.04) |
| $\beta_R$ | $0.45^{***}$ | $0.46^{***}$ | $0.53^{***}$ | $0.35^{***}$ | $0.42^{***}$ | $0.46^{***}$ |
| | (0.15) | (0.15) | (0.11) | (0.11) | (0.12) | (0.13) |
| $\beta_X$ | $-0.10^{**}$ | $-0.11^{**}$ | $0.06^{*}$ | $0.07^{**}$ | | |
| | (0.04) | (0.05) | (0.03) | (0.03) | | |
| $\psi$ | | | | | $0.04^{***}$ | $0.03^{**}$ |
| | | | | | (0.01) | (0.01) |
| $\psi_R$ | | | | | $0.16^{***}$ | $0.19^{***}$ |
| | | | | | (0.03) | (0.04) |
| Factor | Credit | Credit | Market | Market | Corporate | Corporate |
| Classification | History | Contemp. | History | Contemp. | History | Contemp. |
| Rate controls | Yes | Yes | Yes | Yes | Yes | Yes |
| Observations | 8,912 | 5,766 | 8,912 | 5,766 | 12,222 | 8,669 |

*Notes*: Issuer portfolios with issuer fixed effects and the rate controls, interacted with $\mathcal{R}$. Columns (1)–(4) estimate equation (9), in which the export share is the previous year's primary-product share of exports (Clemens and Williamson, 2004), standardized. Columns (5) and (6) estimate equation (10), with a commodity factor that leaves out each borrower's own country; the sample begins in 1872. The Factor row names the factor for $\beta$ and $\beta_R$. Standard errors clustered by four-week period are in parentheses. *** (**) (*) denotes significance at the 1% (5%) (10%) level.

Table 9: Expected losses, carry, and premia by class, 1870–1907

| | Issuers | Default entries | Hazard rate per 1,000 | Haircut % | Expected loss bp | Mean excess return bp | Share of carry % | Premium % a year |
|---|---|---|---|---|---|---|---|---|
| *Panel A: default history to date* | | | | | | | | |
| Safe | 23 | 5 | 0.96 | 49.7 | 4.8 [0.9, 10.7] | 15.4 | 24 [5, 69] | 2.0 [0.2, 3.6] |
| Risky | 21 | 25 | 5.05 | 40.4 | 20.4 [11.7, 31.6] | 35.7 | 36 [21, 63] | 4.6 [1.6, 7.7] |
| *Panel B: contemporary classification* | | | | | | | | |
| Safe | 12 | 2 | 0.53 | 22.6 | 1.2 [0.0, 3.6] | 16.9 | 7 [0, 19] | 2.2 [1.1, 3.5] |
| Intermediate | 7 | 8 | 3.06 | 23.0 | 7.0 [2.1, 14.4] | 28.3 | 20 [6, 71] | 3.7 [0.1, 7.0] |
| Risky | 16 | 20 | 6.29 | 52.8 | 33.2 [17.8, 51.7] | 31.9 | 51 [31, 84] | 4.2 [0.9, 7.5] |

*Notes*: Equation (12), per four-week period, for the 39 sovereign issuers. Issuers: the number with a quoted return in the class; by default history, an issuer moves to the risky class at its first default. Default entries: the 30 entries of Meyer et al. (2022) in issuer-periods with an observed return, classed by the issuer's record before the entry; the 9 entries by issuers with no bond quoted in the panel at the time are excluded. Hazard rate: entries per 1,000 issuer-periods at risk, those with an observed return outside a default episode; the period of a default entry is at risk, so its return enters the mean excess return. Haircut: the mean present-value haircut of the restructurings that follow the class's entries (Meyer et al., 2022). Expected loss: the hazard rate times the haircut. Mean excess return: over the bill, on the issuer-periods at risk. Share of carry: expected loss over implied carry, the mean excess return plus the expected loss. Premium: the mean excess return at an annual rate. Brackets: 95% bands from 2,000 bootstrap draws, with Poisson default entries, haircuts resampled by episode, and class mean returns from a moving-block bootstrap (blocks of 26 periods). Panel B omits the four sovereigns without a contemporary class.

# A Model appendix

This appendix derives the three propositions of Section 3 from the return equation

$$r^{e}_{i,t+1} \;=\; \lambda_t\, q_i \;+\; \beta_i\, u_{t+1} \;+\; d_{i,t+1}, \tag{A.1}$$

the law of motion $\lambda_{t+1} = \bar{\lambda} + \rho(\lambda_t - \bar{\lambda}) - u_{t+1}$ with $0 \le \rho < 1$ and $E_t[u_{t+1}] = 0$, and the two assumptions of Section 3. Under a common price of risk, the same $\lambda_t$ applies to every claim, and the intermediary requires claim $i$ to earn a premium of $\lambda_t q_i$ per period over the bill. Under independent defaults, the default arrival and loss-given-default processes of every claim, with per-period intensity $\pi_i$ and loss given default $\ell_i$, are independent of the initial price of risk and of the sequence $\{u_t\}$ that generates its subsequent path, and so of $\lambda_t$ at every date. The population of issuers whose defaults are counted is fixed, or its inclusion and observation rules are independent of the price-of-risk innovations. The path of the bill rate is not revised by $u$. The additive return equation (A.1) is a first-order specification, with a baseline gross return of one, rather than an exact payoff identity in a period with both default and repricing. Prices and revaluations are in logs; outside default, the difference between log returns and the simple returns of the empirical analysis is second order over a four-week period. The default component and the carry identity in (ii) are in simple returns, in which a default that removes fraction $\ell_i$ of the claim's value lowers the return by $\ell_i$ rather than by $-\log(1-\ell_i)$.

**The price.** Under a common price of risk, the price satisfies the recursion

$$P_{i,t} \;=\; \exp(-r_{f,t} - \lambda_t q_i)\, E_t\big[(1 - L_{i,t+1})(P_{i,t+1} + C_{i,t+1})\big],$$

where $r_{f,t}$ is the bill rate, $C_{i,t+1}$ the promised cash flow, and $L_{i,t+1}$ the realized loss fraction, equal to $\ell_i$ with probability $\pi_i$ and zero otherwise. Claim $i$ promises cash flows at horizons $k = 1, \ldots, T_i$. Let $\omega_{i,k}$ denote the share of the claim's price at date $t$ accounted for by the cash flow at horizon $k$ (the present-value weights, which sum to one), and let $D_i = \sum_k k\,\omega_{i,k}$ be the claim's duration. Iterating the recursion forward, the cash flow at horizon $k$ is discounted, over and above the bill, by the premia expected over the $k$ periods it takes to arrive, $q_i \sum_{j=0}^{k-1} E_t[\lambda_{t+j}]$. To first order in the premium, the log price is thus

$$\log P_{i,t} \;=\; m_{i,t} \;-\; q_i \sum_{k=1}^{T_i} \omega_{i,k} \sum_{j=0}^{k-1} E_t[\lambda_{t+j}], \tag{A.2}$$

where $m_{i,t}$ collects the bill-rate discounting, which is not revised by $u$, and the expected default losses, which do not depend on $u$ under independent defaults. Second-order terms in $q_i\lambda_t$, including the Jensen term in the expectation, are omitted.

**The revaluation coefficient.** By the law of motion, $E_t[\lambda_{t+j}] = \bar{\lambda} + \rho^j(\lambda_t - \bar{\lambda})$, and the innovation $u_{t+1}$ revises $E[\lambda_{t+1+j}]$ between $t$ and $t+1$ by $-\rho^j u_{t+1}$ for every $j \geq 0$. The revision reprices the cash flows outstanding at $t+1$. The cash flow at horizon $k$ from $t$ is then at horizon $k-1$ and is discounted by the premia for the $k-1$ periods still to run, $\lambda_{t+1}, \ldots, \lambda_{t+k-1}$, whose revisions sum to $-u_{t+1}\sum_{j=0}^{k-2}\rho^j$; the cash flow at horizon 1 is paid at $t+1$ and is not repriced. Its weight in the price at $t+1$ is $\omega_{i,k}/(1-\omega_{i,1})$ to zeroth order in the premium, and the price is the share $1-\omega_{i,1}$ of the claim's payoff $P_{i,t+1} + C_{i,t+1}$ to the same order. Multiplying the unexpected change in the log price by that share, the part of the unexpected excess return between $t$ and $t+1$ that $u_{t+1}$ accounts for is

$$q_i\, u_{t+1} \sum_{k=2}^{T_i} \omega_{i,k} \sum_{j=0}^{k-2} \rho^j \;=\; \kappa_i\, q_i\, u_{t+1}, \qquad \kappa_i \;\equiv\; \sum_{k=1}^{T_i} \omega_{i,k}\, \frac{1-\rho^{k-1}}{1-\rho}. \tag{A.3}$$

The coefficient $\kappa_i$ is a $\rho$-discounted duration: the cash flow at horizon $k$ is exposed to the $k-1$ premia not yet earned at $t+1$, each discounted by the persistence of the price of risk. A claim with a single payment at $t+1$ has $\kappa_i = 0$. As $\rho \to 1$, $\kappa_i \to D_i - 1$, the claim's duration measured from $t+1$, with the payment at $t+1$ counted at horizon zero; for $\rho < 1$ it is smaller whenever the claim has cash flows beyond horizon 2; and it is weakly increasing (strictly, for $\rho > 0$) under a shift of weight toward longer horizons. Writing $\Omega_i = \sum_k \omega_{i,k}\rho^{k-1}$, (A.3) gives $\kappa_i = (1-\Omega_i)/(1-\rho) \leq 1/(1-\rho)$, with equality only when $\Omega_i = 0$, that is, when $\rho = 0$ and $\omega_{i,1} = 0$. For a claim whose cash flows lie mostly beyond the horizon over which the price of risk reverts, so that $\Omega_i$ is small, $\kappa_i$ is close to $1/(1-\rho)$ whatever its duration, and the loadings of such bonds are nearly insensitive to maturity. The revaluation term in (A.1) is therefore $\beta_i u_{t+1}$ with

$$\beta_i \;=\; \kappa_i\, q_i. \tag{A.4}$$

Equation (A.4) holds to first order in $q_i\lambda_t$. The weights $\omega_{i,k}$ move with the price, so $\beta_i$ is strictly $\beta_{i,t}$. If the weights are held fixed at those of a consol, $\omega_{i,k} = (1-x)x^{k-1}$ with $0 < x < 1$, then (A.3) gives $\kappa_i = x/(1-x\rho)$.

**Proof of Proposition 1.** Under a common price of risk, the same law of motion, and so the same $\rho$, enters (A.3) for every claim, and under independent defaults, with no bill-rate news in $u$, the term $m_{i,t}$ in (A.2) is not revised by $u_{t+1}$, so (A.3) is the whole of the projection of the unexpected excess return on $u_{t+1}$ and (A.4) gives $\beta_i = \kappa_i q_i$. The coefficient $\kappa_i$ differs across claims only through the cash-flow weights $\omega_{i,k}$. For any two claims with nonzero loadings, $\beta_i/\beta_j = \kappa_i q_i/\kappa_j q_j$; for two claims with $\kappa_i = \kappa_j$, which identical weights imply, $\beta_i/\beta_j = q_i/q_j$, and two such claims with $q_i = q_j$ have $\beta_i = \beta_j$. □

A guarantee of borrower $i$'s debt lowers $q_i$ (its loss covaries less with the rest of the book) and moves $\beta_i$ along the schedule (A.4).

**Proof of Proposition 2.** (i) Let $N_{t+k}$ be the number of default entries at date $t+k$, the sum over claims at risk of indicators with intensities $\pi_i$. For each claim at risk, the probability of a default entry is $\pi_i$ whatever the value of $u_t$, and the set of claims at risk depends only on past defaults and on the population of issuers counted, both independent of $\{u_t\}$ under independent defaults, so $E[N_{t+k} \mid u_t] = E[N_{t+k}]$ for every $k \geq 1$, and the projection of $N_{t+k}$ on $u_t$ has slope zero. The same holds for any issuer-level entry indicator. Independence across issuers is not needed; defaults may be correlated with one another, which is what gives $q_i$ its content, so long as they are independent of $u$.

(ii) Write the default component in (A.1) as $d_{i,t+1} = -(L_{i,t+1} - E_t[L_{i,t+1}])$, so that $E_t[L_{i,t+1}] = \pi_i \ell_i$ and $E_t[d_{i,t+1}] = 0$. Let $s_{i,t}$ denote the claim's promised carry over the bill, the excess return it pays in a period with no default and no unexpected revaluation. Realized excess return is the promised carry less the realized loss plus the revaluation, $r^e_{i,t+1} = s_{i,t} - L_{i,t+1} + \beta_i u_{t+1}$. Taking expectations at $t$, using $E_t[u_{t+1}] = 0$, and matching to (A.1) gives, to first order, the identity

$$s_{i,t} \;=\; \lambda_t \, q_i \;+\; \pi_i \, \ell_i, \tag{A.5}$$

the decomposition of the carry into the premium and the actuarial component; the expected excess return itself is the premium, $E_t[r^e_{i,t+1}] = \lambda_t q_i$. For the revaluation, $\pi_i$ and $\ell_i$ are not functions of $u$ under independent defaults, so the expected default losses in $m_{i,t}$ are not revised by $u_{t+1}$, and with no bill-rate news in $u$ neither is the bill-rate discounting; the projection of the unexpected excess return on $u_{t+1}$ is the term $\kappa_i q_i u_{t+1}$ in (A.3), a revision of the premium path $\{E[\lambda_{t+j}]\}$ alone. The default component $d_{i,t+1}$ is orthogonal to $u_{t+1}$ but is not zero; it is the part of the unexpected return that $u$ does not account for. □

**Proof of Proposition 3.** Under equation (2), the terms $\kappa_i q_i u_{t+1}$ and $d_{i,t+1}$ are as in (A.1)–(A.4), and $\chi_i \zeta_{t+1}$ is a revision of expected payments and default losses that, with $q_i$ held fixed, leaves the premium path unchanged. Abstracting from the rate controls, all covariances below are conditional on time-$t$ information, with variances $\sigma_u^2 > 0$, $\sigma_\nu^2 \geq 0$, and $\sigma_\zeta^2 > 0$.

(i) The loading on $f^m$ alone is

$$\frac{\mathrm{Cov}(r^e_{i,t+1}, f^m_{t+1})}{\mathrm{Var}(f^m_{t+1})} \;=\; \frac{\xi \, \kappa_i q_i \, \sigma_u^2 + \chi_i \, \sigma_\zeta^2}{\xi^2 \sigma_u^2 + \sigma_\zeta^2}, \tag{A.6}$$

which rises with $\chi_i$; the loading on $f^c$ alone, given in (iii), does not involve $\chi_i$.

(ii) The innovation $u_{t+1}$ revises $E[\lambda_{t+j}]$ by $-\rho^{j-1} u_{t+1}$ for $j \geq 1$, so it revises the expected excess return of claim $i$ in period $t+1+j$, which is $\lambda_{t+j} q_i$, by $-q_i \rho^{j-1} u_{t+1}$. Adding the revaluation on impact from (A.3), the cumulative response over periods $t+1, \ldots, t+1+H$ is

$$C_i(H) \;=\; \kappa_i q_i u_{t+1} - q_i u_{t+1} \sum_{j=1}^{H} \rho^{j-1} \;=\; q_i u_{t+1} \Big[ \kappa_i - \frac{1-\rho^H}{1-\rho} \Big], \tag{A.7}$$

to first order, treating the cumulative response as the sum of one-period responses while the claim remains outstanding and $q_i$ is fixed. Using $\kappa_i = (1-\Omega_i)/(1-\rho)$, $C_i(H) = q_i u_{t+1}(\rho^H - \Omega_i)/(1-\rho)$, which converges as $H$ increases to $-q_i u_{t+1}\Omega_i/(1-\rho)$; relative to the impact response, the limit is $\Omega_i/(1-\Omega_i)$ in absolute value, small when $\Omega_i$ is small. For two such claims, the difference in cumulative responses is approximately $(q_r - q_s)\, u_{t+1}\, \rho^H/(1-\rho)$ and decays geometrically at rate $\rho$. The news $\zeta_{t+1}$ leaves expected excess returns unchanged after impact, so its cumulative response is $\chi_i \zeta_{t+1}$ at every horizon.

(iii) The loading on $f^c$ alone is

$$\frac{\mathrm{Cov}(r^e_{i,t+1}, f^c_{t+1})}{\mathrm{Var}(f^c_{t+1})} \;=\; \kappa_i q_i \,\frac{\sigma_u^2}{\sigma_u^2 + \sigma_\nu^2}, \tag{A.8}$$

proportional to $\kappa_i q_i$ with a factor common to every claim. In the projection on $f^c$ and $f^m$ together, with $\mathrm{Var}(f^c) = \sigma_u^2 + \sigma_\nu^2$, $\mathrm{Var}(f^m) = \xi^2\sigma_u^2 + \sigma_\zeta^2$, and $\mathrm{Cov}(f^c, f^m) = \xi\sigma_u^2$, the coefficients are

$$b_i^c \;=\; \frac{\sigma_u^2\sigma_\zeta^2}{|\Sigma_f|}\big(\kappa_i q_i - \xi\chi_i\big), \qquad b_i^m \;=\; \frac{\xi\,\kappa_i q_i\,\sigma_u^2\sigma_\nu^2 + \chi_i\,\sigma_\zeta^2(\sigma_u^2+\sigma_\nu^2)}{|\Sigma_f|}, \tag{A.9}$$

where $|\Sigma_f| = \sigma_u^2\sigma_\zeta^2 + \sigma_\nu^2(\xi^2\sigma_u^2 + \sigma_\zeta^2)$ is the determinant of the covariance matrix of the two factors. The factor $\sigma_u^2\sigma_\zeta^2/|\Sigma_f|$ is common to every claim and falls as $\sigma_\nu^2$ rises. □

# B Data appendix

**Sample.** The sample is the debt of sovereign borrowers serviced in London that carried no metropolitan guarantee. It includes Egypt, Bulgaria, and Romania, which were under Ottoman suzerainty for some or all of the sample, because the suzerain guaranteed none of their loans. The Ottoman loan of 1855, guaranteed by Britain and France, is treated as a backed claim. Hawaii leaves the sample at its annexation by the United States in July 1898, when the United States assumed its public debt. Bonds of British colonies, which carried an implicit metropolitan guarantee (Ferguson and Schularick, 2006; Accominotti et al., 2011), enter only as the comparison group in the test of guarantees.

**Return construction.** A bond's net return over a four-week interval is given by equation (3). It uses the midpoint of bid and ask when both are quoted at both ends of the interval, and otherwise the same quoted side at both ends. No return is formed across a quotation gap, and returns in periods of default are computed from quoted prices like any others.

**The corporate factor.** The index of Chabot and Kurz (2010) covers the domestic British corporate bonds in the December official lists printed in the *Money Market Review*, priced at four-week intervals from January 1866 to December 1907 at the mean of bid and ask prices, with returns including coupons. Securities are classified as domestic by where the company operated, and the index is value-weighted.

**The commodity factor.** The commodity factor is built from the returns of the 1,808 equities quoted on the London Stock Exchange between 1866 and 1907 (Chabot and Kurz, 2010; Alquist et al., 2022), on the same four-week dates as the bonds. The long leg holds the producers of primary commodities operating outside the core (the United Kingdom, France, Germany, the Netherlands and Belgium) and the United States: the oil companies, plantations and mines of that panel's industry classification, and producers of nitrates, meat, sugar, tobacco, timber, coal and iron ore classified elsewhere in it. Gold mines are excluded, because the gold standard fixed the price of their output. The short leg holds the equities of core countries outside the commodity industries. Both legs are weighted by market value at the start of each period, and a leg requires three securities.

**The bank factor.** The British bank factor is the value-weighted return, at lagged market values, on the 51 securities the equity panel classifies as British banks, with at least three in a period, in excess of the bill rate.

**The bill rate.** The banker's bill rate is the open-market discount rate on a banker's bill, a bill of exchange drawn against a shipment or stock of goods and accepted by a London bank. The rate

is quoted as an annual percentage. The quoted bills are short-term London banker's bills with 30–60 days to run, and the annualized rate is geometrically scaled to obtain a four-week return that matches the sampling of the bond prices. When the shortest-maturity bill is unavailable, scaled versions of the two- and three-month bills are used. The data are from *The Economist*.

**The consol yield change.** The consol's current yield is the coupon over the mid quotation. Its four-week change holds the coupon at the previous period's value, so that the coupon reductions of the Goschen conversion, in April 1889 and April 1903, do not enter as changes in yield.

**Digitizing the classifications.** The Powell (1910) classification is the "Locus of Known Borrowers" table (pp. 149–150) with the four "orders" defined in the preceding text (pp. 147–148) and the footnote assigning loans guaranteed by a European government to the first order; the paper's safe, intermediate, and risky classes are Powell's first, second, and third orders. Thirty-five of the 39 issuers appear; Bulgaria, Hawaii, Morocco, and the Orange Free State do not, and they are excluded from estimates that use the contemporary classification. Default episodes are matched from the restructuring record of Meyer et al. (2022) by issuer name.

# C Additional robustness results

Table C.1 reports the fit and volatility of returns by class, Table C.2 the Bank Rate forecast, the crisis exclusions, and the default forecasts, Table C.3 the robustness and backed-claim results, Table C.4 the additional controls and subsamples, and Table C.5 the responses by horizon.

Column (6) of Table 5 adds the sovereign illiquidity factor of Alquist (2010) to the credit-innovation regression, reconstructed for sovereign bonds using that paper's procedure: quotations below 39% of par are dropped, and bonds are sorted each period on their relative bid–ask spread two periods earlier into equal-weighted deciles. The factor is the excess return of the widest-spread decile over the narrowest.

Table C.1: Fit and volatility of sovereign returns, by borrower class

| | Default history | | Contemporary | |
|---|---|---|---|---|
| | Safe | Risky | Safe | Risky |
| Loading on the corporate factor | 0.15 | 0.58 | 0.15 | 0.62 |
| Within $R^2$ | 0.042 | 0.021 | 0.047 | 0.018 |
| SD of excess return, % per period | 2.23 | 7.27 | 2.05 | 8.12 |
| Residual SD, % per period | 2.18 | 7.19 | 2.00 | 8.05 |
| Issuer-periods | 5,181 | 7,467 | 3,819 | 5,108 |

*Notes*: Each class's issuer portfolios are regressed separately on the corporate factor and the rate controls with issuer fixed effects. Within $R^2$ is the share of the within-issuer variance explained. The standard deviations are of within-issuer four-week excess returns and residuals.

Table C.2: Bank Rate forecast, crisis periods, and default forecasts

*Panel A: forecast of the Bank Rate change, equation (6)*

| $\tau_1$ | $\tau_2$ | $\tau_3$ | $R^2$ | Periods |
|---|---|---|---|---|
| $0.40^{***}$ | $0.19^{***}$ | $-0.24^{***}$ | 0.20 | 494 |
| (0.07) | (0.05) | (0.03) | | |

*Panel B: Bank Rate pass-through with crisis periods excluded*

| | $\gamma_S$ | $\gamma_I$ | $\gamma_R$ | $\gamma_R - \gamma_S$ |
|---|---|---|---|---|
| History | $-0.65^{***}$ | | $-1.18^{***}$ | $-0.53^{**}$ |
| | (0.12) | | (0.23) | (0.25) |
| Contemp. | $-0.55^{***}$ | $-0.79^{***}$ | $-1.46^{***}$ | $-0.90^{***}$ |
| | (0.09) | (0.16) | (0.27) | (0.25) |

*Panel C: Poisson forecasts of default entries, equation (11)*

| | Credit innovation | | | | | Market innovation |
|---|---|---|---|---|---|---|
| | 1 year | 2 years | 3 years | 4 years | 5 years | 1 year |
| $\vartheta_h$ | $-0.040$ | $-0.057$ | $-0.037$ | $-0.024$ | $-0.021$ | $0.001$ |
| | (0.049) | (0.043) | (0.032) | (0.025) | (0.020) | (0.033) |
| Periods | 480 | 467 | 454 | 441 | 428 | 480 |

*Notes*: Panel A: the forecast regression (6) of the four-week change in Bank Rate, with Newey–West standard errors. Panel B: equation (7), with the year after each chronology crisis (1873, 1878, 1890) and the periods from October 1907 dropped from both the forecast and the response regressions, and moving-block bootstrap standard errors. Panel C: equation (11), the Poisson regression of default entries over the next one to five years on the innovation, with Bartlett standard errors. *** (**) (*) denotes significance at the 1% (5%) (10%) level.

Table C.3: Robustness of the difference in loadings on the corporate factor

Dependent variable: excess return over the bill rate, $r^e_{i,t}$

| | Weighting and gaps | | | Periods dropped | | | Risky split | | Backed claims | | One issuer dropped | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) | (11) | (12) |
| $\beta$ | $0.15^{***}$ | $0.14^{***}$ | $0.14^{***}$ | $0.14^{***}$ | $0.13^{***}$ | $0.15^{***}$ | $0.16^{***}$ | $0.15^{***}$ | $0.14^{***}$ | $0.86^{***}$ | | |
| | (0.04) | (0.04) | (0.04) | (0.04) | (0.04) | (0.04) | (0.04) | (0.04) | (0.04) | (0.32) | | |
| $\beta_R$ | $0.43^{***}$ | $0.51^{***}$ | 0.25 | $0.45^{***}$ | $0.38^{**}$ | $0.46^{***}$ | | | | | $0.43^{***}$ | $0.52^{***}$ |
| | (0.12) | (0.14) | (0.16) | (0.13) | (0.15) | (0.14) | | | | | (0.13) | (0.15) |
| $\beta_{R,P}$ | | | | | | | $0.24^{*}$ | $0.29^{**}$ | | | | |
| | | | | | | | (0.13) | (0.15) | | | | |
| $\beta_{R,D}$ | | | | | | | $0.93^{***}$ | $0.84^{***}$ | | | | |
| | | | | | | | (0.24) | (0.23) | | | | |
| $\beta_B$ | | | | | | | | | 0.02 | $-0.75^{**}$ | | |
| | | | | | | | | | (0.04) | (0.32) | | |
| Classification | Contemp. | Contemp. | Contemp. | Contemp. | Contemp. | Contemp. | History | Contemp. | — | — | Contemp. | Contemp. |
| Unit | Issuer | Bond | Bond | Issuer | Issuer | Issuer | Issuer | Issuer | Bond | Bond | Issuer | Issuer |
| Issuer dropped | | | | | | | | | | | Spain | China |
| Observations | 8,927 | 22,675 | 22,675 | 8,216 | 7,463 | 8,435 | 12,648 | 8,927 | 26,883 | 2,823 | | |

*Notes*: Equation (5) with issuer or bond fixed effects and the rate controls, interacted with $\mathcal{R}$. Column (1) uses equal-weighted issuer portfolios. Columns (2) and (3) use bonds, carrying the last quotation through gaps (2) or assigning the realized return across a gap to its first missing period (3). Column (4) drops the year after each chronology crisis (1873, 1878, 1890) and the periods from October 1907; columns (5) and (6) drop the four-week periods in which Bank Rate stood at or above 5% and 6% on any day. Columns (7) and (8) split the risky class into issuer-periods outside ($\beta_{R,P}$) and inside ($\beta_{R,D}$) a default episode (Meyer et al., 2022). Columns (9) and (10) interact the factor with *Backed* ($\beta_B$), so that $\beta$ is the loading of the claims without backing: colonial bonds beside safe sovereign bonds in (9), and within Turkey the loan guaranteed by Britain and France in 1855 beside the 15 issues without a guarantee or a pledge of the Egyptian tribute in (10). Columns (11) and (12) report the smallest and largest difference over the 28 regressions that each drop one safe or risky issuer. Standard errors clustered by four-week period are in parentheses. *** (**) (*) denotes significance at the 1% (5%) (10%) level.

Table C.4: Additional controls and subsamples for the difference in loadings

Dependent variable: excess return over the bill rate, $r^e_{i,t}$

| | Constraints on the executive | | | GDP per capita | | Gold standard | | | Baring crisis | | Default entries | | | Illiquidity |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) | (11) | (12) | (13) | (14) |
| $\beta$ | 0.19*** | 0.14* | 0.18*** | 0.18*** | 0.15** | | | | | | 0.16*** | 0.16*** | | 0.12*** |
| | (0.05) | (0.09) | (0.05) | (0.05) | (0.06) | | | | | | (0.04) | (0.04) | | (0.04) |
| $\beta_R$ | 0.37*** | 0.49*** | 0.39*** | 0.40*** | 0.47*** | 0.31*** | 0.26** | 0.23* | | | 0.50*** | 0.42*** | 0.47*** | 0.40*** |
| | (0.12) | (0.18) | (0.13) | (0.14) | (0.18) | (0.10) | (0.11) | (0.12) | | | (0.14) | (0.14) | (0.13) | (0.13) |
| $\beta_C$ | −0.06 | 0.01 | −0.06 | | | | | | | | | | | |
| | (0.04) | (0.09) | (0.05) | | | | | | | | | | | |
| $\beta_Y$ | | | | −0.05 | −0.01 | | | 0.04 | | | | | | |
| | | | | (0.05) | (0.08) | | | (0.04) | | | | | | |
| $\beta_G$ | | | | | | −0.30*** | −0.34*** | −0.12 | | | | | | |
| | | | | | | (0.08) | (0.10) | (0.10) | | | | | | |
| $\beta_D$ | | | | | | | | 0.54** | | | | | | |
| | | | | | | | | (0.21) | | | | | | |
| $\delta_N$ | | | | | | | | | | | | | 0.00 | |
| | | | | | | | | | | | | | (0.00) | |
| $\beta^{\text{il}}$ | | | | | | | | | | | | | | 0.07** |
| | | | | | | | | | | | | | | (0.03) |
| $\beta^{\text{il}}_R$ | | | | | | | | | | | | | | 0.14** |
| | | | | | | | | | | | | | | (0.07) |
| $(\beta+\beta_R)/\beta$ | | | | | | | | | 3.0 | 4.4 | | | | |
| | | | | | | | | | (0.9) | (1.8) | | | | |
| Factor | Corporate | Corporate | Corporate | Corporate | Corporate | Corporate | Corporate | Corporate | Corporate | Corporate | Corporate | Corporate | Corporate | Credit |
| Classification | History | Contemp. | History | History | Contemp. | History | Contemp. | History | Contemp. | Contemp. | Contemp. | Contemp. | Contemp. | Contemp. |
| Sample or measure | 1870–1907 | 1870–1907 | 1850–69 | 1865–70 | 1865–70 | Full | Full | Prev. year | Before | After | Entries out | With next | Full | Full |
| Factor × decade | No | No | No | No | No | No | No | Yes | No | No | No | No | No | No |
| Observations | 12,297 | 8,927 | 12,079 | 12,379 | 8,927 | 12,648 | 8,927 | 12,379 | | | 8,239 | 7,640 | 8,927 | 8,825 |

*Notes*: Equation (5) on issuer portfolios with issuer fixed effects and the rate controls, interacted with $\mathcal{R}$; the Factor row names the factor for $\beta$ and $\beta_R$. The factor is interacted with a standardized composite of V-Dem's judicial and legislative constraints on the executive ($\beta_C$), averaged over 1870–1907 or over 1850–69 (Coppedge et al., 2026), in columns (1)–(3); with standardized log GDP per capita in 1865–70 ($\beta_Y$) in columns (4) and (5); and with an indicator for adherence to the gold standard in the period ($\beta_G$) in columns (6) and (7). Column (8), on the 37 sovereigns with GDP data, also interacts it with decade, with an indicator for a default episode ($\beta_D$), and with standardized log GDP per capita in the previous year ($\beta_Y$). Columns (9) and (10) report $(\beta+\beta_R)/\beta$ before and after the Baring crisis of November 1890, with delta-method standard errors. Column (11) drops the four-week periods that contain a default entry, column (12) also drops the period after each, and column (13) adds the number of entries in the period ($\delta_N$). Column (14) adds the sovereign illiquidity factor ($\beta^{\text{il}}$, $\beta^{\text{il}}_R$) to the credit-innovation regression. Standard errors clustered by four-week period are in parentheses. *** (**) (*) denotes significance at the 1% (5%) (10%) level.

Table C.5: Responses by horizon

*Panel A: cumulative excess return by class, local projections*

| | Credit innovation | | Market innovation | |
|---|---|---|---|---|
| | Impact | One year | Impact | One year |
| | (1) | (2) | (3) | (4) |
| Safe | $0.17^{**}$ | $0.57^{**}$ | $0.28^{***}$ | $0.37^{*}$ |
| | (0.07) | (0.23) | (0.06) | (0.19) |
| Intermediate | $0.31^{**}$ | $0.35$ | $0.57^{***}$ | $0.61^{*}$ |
| | (0.13) | (0.47) | (0.09) | (0.32) |
| Risky | $0.48^{***}$ | $0.35$ | $0.95^{***}$ | $1.39^{***}$ |
| | (0.16) | (0.59) | (0.18) | (0.33) |
| Risky − safe | $0.31^{***}$ | $-0.22$ | $0.68^{***}$ | $1.02^{***}$ |
| | (0.12) | (0.49) | (0.14) | (0.27) |

*Panel B: export growth in a single year, default history to date*

| | Credit innovation | | | Market innovation | | |
|---|---|---|---|---|---|---|
| | Year $s$ | Year $s+1$ | Year $s+2$ | Year $s$ | Year $s+1$ | Year $s+2$ |
| | (5) | (6) | (7) | (8) | (9) | (10) |
| $\eta$ | $-0.12$ | $-0.30$ | $-0.52$ | $2.04^{***}$ | $2.89^{***}$ | $0.98^{*}$ |
| | (0.67) | (0.81) | (0.39) | (0.41) | (0.62) | (0.55) |
| $\eta_R$ | $0.42$ | $-0.39$ | $-0.66$ | $0.57$ | $-0.19$ | $-0.89$ |
| | (1.14) | (0.74) | (0.74) | (1.00) | (1.42) | (0.88) |
| Issuer-years | 712 | 733 | 739 | 712 | 733 | 739 |

*Notes*: Panel A: local projections of the mean cumulative excess return of each class's bonds, under the contemporary classification (Powell, 1910), on the credit or market innovation, at impact ($h = 0$) and at one year ($h = 13$), with Newey–West standard errors of lag $h + 1$. Panel B: equation (8) with single-year export growth as the dependent variable, in percentage points per one-standard-deviation annual innovation, with Driscoll–Kraay standard errors. *** (**) (*) denotes significance at the 1% (5%) (10%) level.